Labor Composition and U.S. Productivity Growth, 1948-90

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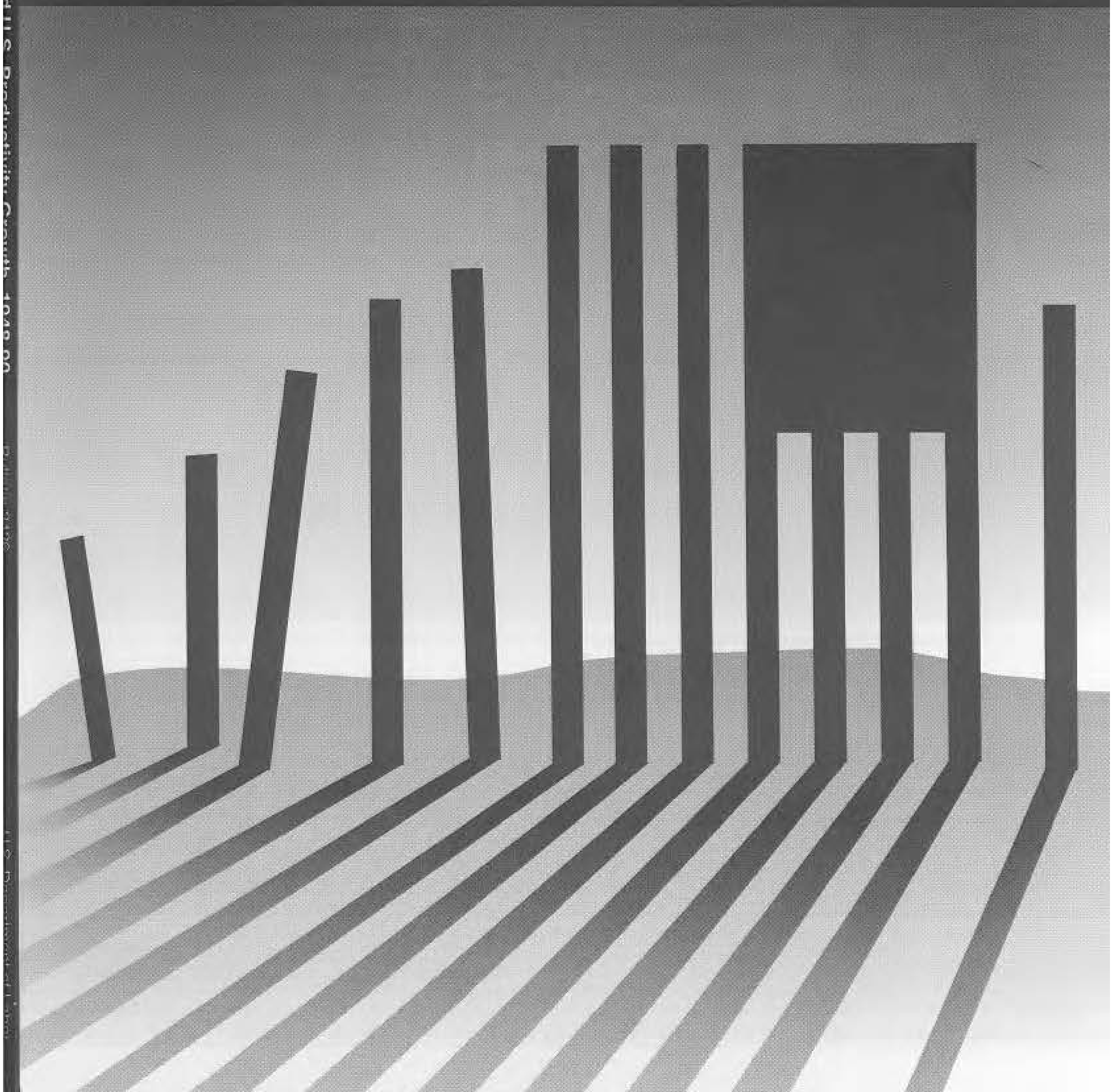

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U.S. Department of Labor
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Preface

A principal function of the Bureau of Labor Statistics (BLS) is to inform policymakers and the general public on factors influencing the well-being of U.S. workers. Thus an important part of the Bureau's work is the study of productivity, which relates directly to workers' real incomes, price stability, employment, and the competitiveness of U.S. goods and services in world markets.

The major purpose of this bulletin is to present new BLS measures of changes in the composition of the labor force that have influenced productivity growth. The bulletin also provides an estimate of how much changes in labor force composition contributed to the growth in labor productivity (output per hour) for the years 1948 to 1990. The bulletin recognizes that the effective quantity of labor input does not rest solely on the total number of hours worked by members of the U.S. labor force but also on characteristics of the labor force. This study focuses on changes in the distribution of hours worked by educational attainment and years of work experience for both men and women. The study also examines the role of labor force composition change in the important slowdown in productivity growth that began about 1973 and that has hampered the growth in U.S. living standards since then.

The bulletin provides measures of the total labor input used in the production of output in the private business sector and the private nonfarm business sector of the U.S. economy from 1948 to 1990. The trend in total labor input divides into the trends in hours worked and in labor force composition. The trend in labor productivity—output per hour worked—is related not only to the trend in labor composition but also to trends in capital input and in multifactor productivity.

Few people doubt that education and training have an important influence on a nation's output and productivity growth and consequently on the well-being of its citizens. Recently, much public discussion has been devoted to the link between education and training and productivity growth. Numerous researchers and policymakers have examined ways to stimulate firms to improve education and job skills through formal and informal on-the-job training in technologically advanced occupations. This study does not address all facets of the recent public discussion. It does examine, however, the past record of how educational attainment and work experience contributed to the growth in productivity. The study's results may therefore be relevant to some aspects of the public discussion on productivity growth.

BLS measures labor composition change in keeping with a recommendation of the Panel to Review Productivity Statistics set up by the National Academy of Sciences and chaired by the late Professor Albert Rees. In its 1979 report, *Measurement and Interpretation of Productivity*, the panel recommended that

> . . . the Bureau of Labor Statistics devote more resources to studying the use of weighted labor input measures. The purpose of this effort would be the preparation of one or more weighted measures of labor input for broad aggregates of economic activity, such as the private business sector. (p. 128)

The panel set forth a number of reasons for developing such weighted measures and also examined potential labor force characteristics that should be used in developing the measures. It suggested the "use of weights based on some skill-related dimension of labor input, such as years of schooling, formal on-the-job training, or years of experience" (p. 127). In several respects, the present study fulfills these recommendations.

Chapters I, II, and III present the main results of the Bureau's study, and several technical appendixes present the Bureau's methods, data sources used, sensitivity tests designed to show the effects of alternative methodological approaches on the main results, and comparisons of the Bureau's approach to labor composition measurement with the approaches of other studies.

The measures published in this bulletin are the latest in a series of Bureau measures designed to identify the sources of change in labor productivity. In 1983, the Bureau presented its measures of capital input and multifactor productivity in Bulletin 2178, *Trends in Multifactor Productivity*. These capital input and multifactor productivity measures have been periodically updated since they were introduced. In 1989, the Bureau presented measures of the direct effects of expenditures on research and development on productivity change, in Bulletin 2331, *The Impact of Research and Development on Productivity Growth*.

In other recent publications, the Bureau has analyzed multifactor productivity for detailed industries using gross output rather than value added and including purchased intermediate inputs in addition to the inputs of labor and capital. BLS has also examined, in a number of studies, alternative methods of measuring capital inputs; the results of these studies have been used to improve the 1983 measures of capital input.

Additional studies have attempted to improve the methods used to measure trends in labor productivity and multifactor productivity or to explain changes in productivity trends, such as the slowdown in productivity growth that occurred after 1973. These studies have focussed on such subjects as the measurement of capacity utilization, the role of rising energy prices in the productivity slowdown, the development of measures of labor input as hours worked—the preferred measure—in place of hours-paid measures, the consequences of the possible failure of the equilibrium assumptions adopted in most approaches to multifactor productivity measurement, and the development of multifactor productivity measures for other major industrialized economies. Most of these post-1983 efforts have been devoted to research issues identified in 1979 and recommended for further study by the National Academy of Sciences panel.

The present study was prepared by the Bureau's Office of Productivity and Technology, under the general direction of Edwin R. Dean, Associate Commissioner. Michael J. Harper, Chief of the Division of Productivity Research, directly supervised the research staff. Larry Rosenblum wrote most of the text of the report, made important contributions to the development of the study, and has been directly in charge of the study in recent years. Kent Kunze directed the study for several years, made important contributions to its development, and organized and drafted sections of the study. Mary Jablonski participated in the research, wrote drafts of several chapters and appendixes, and made other highly valued contributions to the study. The specific research approach adopted in the study was originated by the late William Waldorf, former Chief of the Productivity Research Division. Linda Moeller has provided keen insight into future directions for this work. Professor Michael Tannen, of the University of the District of Columbia, worked on the study under a contract and made important contributions to the development of the study.

The approach to this study and preliminary empirical results have been presented at a number of conferences of economists and other interested experts. In addition, papers reporting preliminary results of the study have been published in conference proceedings volumes and in journals. Finally, in meetings and in correspondence, the approach used in the study has been discussed with other researchers. The comments received through all these channels have been used to refine the approach and methods used in this report and the Bureau is grateful to those who have provided comments. Special thanks are due to Dale Jorgenson, Harvard University; Jack Triplett, Bureau of Economic Analysis, U.S. Department of Commerce; Martin Baily, University of Maryland; Jacob Mincer, Columbia University; Kevin Murphy, University of Chicago; and John Kendrick, George Washington University. Many fruitful discussions were held with the late Edward Denison, of the Brookings Institution. Jerome Mark, former Associate Commissioner of the Bureau's Office of Productivity and Technology, Marilyn Manser, Assistant Commissioner of the Bureau's Office of Economic Research, and Mark Loewenstein, also of the Office of Economic Research, provided useful comments during the course of the research.

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Chapter I. Introduction

This bulletin presents productivity measures for the private business and private nonfarm business sectors of the U.S. economy which reflect new and more comprehensive measures of labor input. Defined generally, productivity is a comparison of outputs and inputs. Traditionally, the Bureau of Labor Statistics (BLS) has defined productivity as output per hour and publishes a wide array of such measures.[1] In September 1983, BLS published *Trends in Multifactor Productivity, 1948-81* (Bulletin 2178) which introduced measures of multifactor productivity. This productivity measure compares output to inputs of both labor and capital and helps explain the effects of changes in capital services on output per hour. The present bulletin extends earlier research, focusing on a newer concept of labor input.

Output per hour measures focus on the time dimension of labor input measures. Labor input is based mainly on the employment and average weekly hours of employees from the Current Employment Statistics (CES) survey of nonagricultural establishments and the number and hours of self-employed persons, unpaid family workers, and agricultural workers from the Current Population Survey (CPS). Accordingly, aggregate labor input is defined as the sum of the hours of these workers—differences in the effectiveness of an hour of labor input across workers were not considered.

The effectiveness of labor, however, may change because of changes in worker effort, the organization of the workplace, or the composition of the work force (as well as changes in nonlabor inputs). Previous efforts to account for output per hour change are extended in this study by measuring one aspect of change in the effective quantity of an average hour of work: changes in the composition of the work force. The effective quantity of labor input depends not only on the hours of labor but also on the characteristics of those performing the work. Consequently, workers of differing traits are not perfect substitutes, and firms treat them as distinct inputs in their production process. In this bulletin, the hours of workers are classified according to educational attainment and work experience for each sex. By weighting the changes in the hours of each type of labor, changes in the educational attainment and experience of the work force are incorporated into the measures of aggregate labor input for the period 1948 through 1990. It is then possible to measure the impact of changes in labor force composition on productivity growth during this period.

The remainder of this chapter provides an overview of how changing work force characteristics affect measures of labor input and productivity and how these changes can be incorporated into the productivity measures. First, concepts and definitions of productivity, including labor productivity and multifactor productivity, are discussed. The concepts of weighted labor input and labor composition are then presented and related to measures of productivity. After a brief description of the data used in this study, measures of labor composition change and its effect on multifactor productivity growth in the private business and private nonfarm business sectors are summarized. The chapter concludes by describing the organization of the bulletin.

Productivity

Productivity is defined as the ratio of output to one or more inputs. Productivity increases if output grows more than inputs. Conversely, productivity declines if output grows less than inputs. Economic growth can be achieved through increases in the quantity of inputs or productivity growth. Increases in labor, capital, and material inputs impose costs on society such as less leisure time, lower current consumption (because of increased investment), and lower reserves of natural resources. Productivity growth represents increases in output not attributable to increases in factor inputs. Therefore, productivity growth is the primary measure of changes in the efficiency of the economy. Without productivity growth there are few means of permanently raising the well-being of the average worker. Given finite resources, a lack of productivity growth would mean that economic changes would largely reflect shifts in the distribution of resources and wealth across the population.

Productivity growth occurs through a number of conduits. Growth can arise from new technology, enhanced worker motivation, economies of scale in production, improved managerial skills (which result in better use of resources), increased worker skills (which result in better use of the available technology and capital services), and other sources.

Measures of productivity

The most common measure of productivity is output per hour or "labor productivity." This measure might be referred to as a partial productivity measure because it is based on only one of several factor inputs. Changes in labor productivity reflect the growth in output not attributed to the growth in labor. Therefore, labor productivity changes in response to all the

[1] BLS publishes quarterly measures of output per hour for the business, nonfarm business, and manufacturing sectors. Annual measures for more than 200 detailed industries as well as international comparisons of manufacturing productivity are also available. A summary of the BLS productivity program can be found in chapters 10-12 of the *BLS Handbook of Methods* (Bulletin 2414).

sources already mentioned *and* to changes in the quantities of factor inputs other than labor.

For example, a farmer who buys a new tractor that can cultivate 500 acres of land instead of 200 acres during the growing season will produce more output. An increase in labor productivity will be measured. This increase is clearly attributable to increased capital services (a new and faster tractor). If capital services are, in practice, not measured, an increase in measured labor productivity could be due either to changes in the efficiency of the farm or to increased capital services. Labor productivity measures, then, do not distinguish between changes in the efficiency of the farm and changes in other inputs. The advantage of labor productivity over more complex productivity measures is that the necessary data (measures of labor and output) are more readily available and are available sooner. Thus the measurement process is much more convenient, and the measures are more timely.

Multifactor productivity (MFP) is a more comprehensive measure of productivity which is not subject to changes in unmeasured inputs. Changes in multifactor productivity are defined as the growth rate of output less the growth rate of aggregate input, which is measured by the growth rate of each input weighted by its share of the total cost of production. Because these share weights sum to one, the change in the aggregate inputs is a weighted average of the growth rates of each input. Changes in multifactor productivity can be viewed as the growth rate of output less a weighted average of the growth rates of the inputs. A formal derivation of the growth accounting framework used to measure multifactor productivity can be found in appendix A.

For major sectors of the economy, such as the private business sector, most productivity analysts are largely concerned with how efficiently the economy satisfies the needs of individuals for consumption goods and of firms for capital goods. Some outputs such as energy, materials, or business services are largely sold to other firms to be used as inputs by these firms. The value of such inputs adds to the output of each firm in the chain to the final consumer, and the net sale (or output) to consumers of these intermediate inputs is near zero. These "intermediate outputs" are excluded from both the measures of output and inputs to avoid double counting and possible bias in productivity measures.[2]

Labor productivity and multifactor productivity are closely related concepts. At the aggregate level, the growth rate of output per hour (labor productivity) is the sum of the growth rates of multifactor productivity and the effects of capital intensity (defined as capital services per hour weighted by capital's share of total costs). Consequently, multifactor productivity measures the effects of changes in new technology, enhanced worker motivation, economies of scale in production, improved managerial skills (which result in better use of resources), increased worker skills (which result in better use of the available technology and capital services), and other sources. Multifactor productivity does not measure changes in capital intensity.

Labor input measurement

Labor input reflects the time, effort, and skills of the work force. Previous BLS measures of productivity change have focused exclusively on the time dimension of labor input. That is, most measures of labor input are simply the sum of all hours of workers. However, workers have differing traits and are usually not interchangeable. The fact that firms pay workers with different characteristics at different rates shows that firms do not regard all workers as identical. Accordingly, workers with differing traits should be treated as separate and distinct inputs in the measurement of output and productivity changes.

In a comprehensive study of the methodology of productivity measurement, the National Academy of Sciences recommended that differences among workers should be incorporated into the measures of labor input.[3] Recent studies by a number of economists have focused on how worker skills affect measures of labor input.[4] In *Trends in Multifactor Productivity, 1948-81* (1983), the Bureau of Labor Statistics noted that changing labor input usually represents a change in both the skill level of the work force and aggregate hours. Consequently, the change in the composition of the work force is a source of output growth which can add to our understanding of changes in output per hour.

This study broadens the concept of labor input to reflect differences in the skills of workers. Although there are many dimensions to a worker's skills, this study is limited to two important measures of skills: education and work experience. The hours of workers are classified according to their level of skill as measured by their education and work experience for each sex. (For brevity, the term "skills" is used synonymously with education and experience.)

From the perspective of formal production theory, it might be thought desirable to classify workers by as many relevant traits as possible. Presumably that would yield a good approximation to the "true" labor composition measure. Jorgenson, Gollop, and Fraumeni (1987) have estimated labor composition using a very large number of categories representing a cross-classification of five characteristics (age, education, class of worker, occupation, and sex) for each industry. From the perspective of production theory, this has the advantage

[2] Intermediates are included in both output and input for BLS productivity measures for detailed industries. For a discussion, see William Gullickson and Michael Harper, "Multifactor Productivity in U.S. Manufacturing, 1949-83," *Monthly Labor Review*, October 1987.

[3] *Measurement and Interpretation of Productivity*, Committee on National Statistics, Assembly of Behavioral and Social Sciences, National Research Council, (National Academy of Sciences, Washington, DC, 1979). See recommendation no. 9, p. 128.

[4] Studies of labor force composition include Edward Denison, *Trends in American Economic Growth 1929-1982*, Brookings Institution, (Washington, DC, 1985); D. Jorgenson, F. Gollop, and B. Fraumeni, *Productivity and Economic Growth*, Harvard University Press, (Cambridge, MA, 1987); and Peter Chinloy, "Sources of Quality Changes in Labor Input," *American Economic Review*, March 1980, pp. 108-119. See appendix G for a discussion of the theoretical and empirical differences between these studies and the labor composition measures developed in this bulletin.

that labor composition will reflect not only the direct contribution of the specified characteristics, but may also include the many effects of correlated traits not specifically included within the framework.

Such a comprehensive set of traits, however, can lead to difficulties in identifying sources of labor composition growth. For example, shifts in the proportion of part-time employment may be a source of labor composition change even though work schedules are not a classifying characteristic.

This study narrows its focus in order to identify and measure the effects on productivity of a specific known set of factors. Thus, rather than partition MFP growth into two catch-all categories ("labor composition effects" and "all other sources of multifactor productivity growth"), a labor composition measure is developed that attempts to identify the separate and interacted effects on productivity of changes in two specific characteristics: education and work experience. Because earnings of men and women differ markedly, the method is applied separately to each.

To measure labor input, changes in the distribution of the hours of workers are weighted by their contribution to output. Although one cannot directly observe a worker's contribution to output, in competitive markets, a worker's hourly contribution can be equated to his or her hourly compensation.[5] Differences in hourly compensation rates by the level of education and work experience have been thoroughly documented in hundreds of studies.[6]

The change in weighted labor input is defined as the weighted change in the hours of each type of worker. The weights are the workers' shares of labor compensation which are determined by the relative size and hourly compensation of the group of workers.[7] Consequently, weighted labor input now reflects the skill distribution of the work force, and multifactor productivity can be measured net of the effects of changing work force composition.

In contrast, previous work defined labor input as the sum of all hours of all workers. This can also be viewed as a weighted labor input where the weights for all types of labor are identical. Identical weights imply that all workers contribute equally to production and differences between workers do not matter. Therefore, previous measures of labor input did not reflect the skill distribution of the work force, and as a result, productivity included the effect of changing work force composition.

Of primary interest are measures of labor composition. The change in labor composition is defined as the difference between the growth rate of the weighted labor input and the growth rate of total hours. Increases in labor composition represent a rising average level of worker skills as measured by education and work experience for each sex. By separating the growth of labor input into two components, total hours and labor composition, the contribution of the changing levels of educational attainment and work experience to productivity and economic growth can be analyzed. The contribution of labor composition to output and labor productivity is the product of its growth rate and labor's share of total costs.

Data sources

This study draws on a wealth of data sources. The hours of workers are collected from the annual demographic file of the Current Population Survey for the period 1968 to 1990 and from the decennial censuses of 1950 and 1960. Using these and additional data sources, a matrix of hours is constructed for each year from 1948 to 1990. Initially the matrix is cross-classified by 7 levels of education and 72 age levels for men. A more detailed matrix for women includes the same education and age groups as well as two categories for marital status and four categories for number of children.

A matched sample of the 1973 Current Population Survey, Social Security work histories, and Internal Revenue Service records makes it possible to estimate work experience for each group of workers based on their demographic characteristics. The matrix of hours can then be collapsed into 7 education and 72 work experience levels for each sex.

Hourly wage rates for each category of hours are also constructed annually based on the decennial censuses and Current Population Surveys. The effects of education and work experience on the hourly wage rate of a sample of survey respondents is econometrically estimated. Separate estimates are made for each year. The estimated parameters are used to develop measures of hourly earnings for each of 1,008 types of workers and to aid in examining the separate contribution of education and work experience to productivity growth.

Econometrically estimated wage rates are quite common, but their use in labor composition studies is rare.[8] A wage rate for each type of worker could have been based on the sample average of the Current Population Survey. Changes in the mean wage of a set of workers reflect changes in observed and unobserved characteristics. For example, the average wage for a set of workers may change not because the reward to education or experience has changed but simply because the proportion of part-time workers has shifted. Other characteristics of the work force could have been added to the study and resulted in a broader measure of labor composition change.

[5] This result follows under the assumption that firms face production functions characterized by constant returns to scale and competitive input and product markets. It is also assumed that firms maximize their profits by exactly equating a worker's hourly contribution to his or her hourly compensation. See appendix A for further discussion of this point and for how deviations from competition may affect the relationship between a worker's wage and contribution to output.

[6] The work of Gary Becker, *Human Capital*, Columbia University Press, (New York, NY, 1964) and Jacob Mincer, *Schooling, Experience and Earnings*, Columbia University Press, (New York, NY, 1974) has spawned literally hundreds if not thousands of studies on why and how much education and experience affect earnings.

[7] A Tornqvist index is used to combine the growth rates of the hours of each type of worker into a composite growth rate of labor input. A Tornqvist index effectively weights the growth rate of the hours of each group of workers by their share of labor compensation. Further discussion of Tornqvist indexes can be found in appendix A.

[8] Perhaps the only other study is Steven Allen, "Unionized Construction Workers are More Productive," *Quarterly Journal of Economics*, May 1984, pp. 251-274.

but the sources of labor composition change would have been less clear. Consistent with this more narrow measure of labor composition is an econometrically estimated wage model which permits earnings differences arising from education and work experience to be isolated from other sources of earnings differentials.

The use of work experience rather than age or years since leaving school is an important contribution of this study. Research by Mincer, Becker, and others indicates that the apparent earnings variation by age largely reflects differences in the amount of on-the-job training. Training relates more closely to work experience than to age, thus work experience is a more appropriate dimension for identifying differences between workers. More to the point, the distribution of hours (and to some extent hourly compensation) by work experience differs from the distribution by age, so classification by work experience will affect the measures of labor input and composition.

Although using the matched sample to estimate work experience strengthens the conceptual basis of this study, only a single matched sample is available. Rising labor force participation of women, changing incentives to retire, and delayed entry of youth into the work force due to longer periods of schooling may alter the rate at which workers gain work experience over time. An experience model estimated at a single point in time may fail to capture these significant changes. This possibility is examined in Appendix F, and the labor composition measures are found to be largely unaffected by the use of a single experience model.

Summary of findings

Table 1 presents trends in two important characteristics of the work force over the last 40 years. A steady and persistent increase in the average level of educational attainment is the dominant trend in the work force. Educational attainment increased for both men and women, and for both the increase was at least 3 years. The level of work experience rose between 1948 and 1958 and then proceeded to decline in an irregular and modest fashion until about 1979. Since then, the level of work experience for men has remained more or less constant, while the work experience of women has increased about half a year. For the entire period of this study, the decline for men was greater than for women. The level of work experience is estimated to decline 1 year for men and negligibly for women.

The effects of these two trends in educational attainment and work experience (as well as trends in the relative compensation of workers) are combined to develop the Tornqvist index of labor input presented in table 2. For private business, weighted labor input grew at an annual average rate of 1.3 percent between 1948 and 1990. Because hours worked grew 1.0 percent annually over the same period, the index of labor composition grew 0.3 percent annually. To measure the contribution to productivity growth, labor composition changes are weighted by labor's share of total costs (about two-thirds).

Table 1. Average years of completed schooling and experience of working men and women in private business[1], 1948-90

Year	Education		Experience	
	Men	Women	Men	Women
1948	9.3	9.9	18.8	12.4
1949	9.2	10.2	19.0	12.5
1950	9.2	10.2	19.1	12.5
1951	9.3	10.1	19.3	12.7
1952	9.4	10.1	19.4	12.9
1953	9.5	10.2	19.6	13.0
1954	9.7	10.2	19.8	13.1
1955	9.8	10.3	19.7	13.3
1956	9.9	10.4	19.8	13.5
1957	10.0	10.5	19.8	13.6
1958	10.1	10.5	19.9	13.7
1959	10.3	10.6	19.8	13.7
1960	10.2	10.5	19.8	13.4
1961	10.4	10.7	19.8	13.6
1962	10.6	10.8	19.8	13.5
1963	10.7	10.8	19.8	13.6
1964	10.8	10.9	19.7	13.6
1965	10.9	10.9	19.6	13.5
1966	10.9	11.0	19.6	13.4
1967	11.1	11.0	19.6	13.2
1968	11.2	11.2	19.6	13.1
1969	11.4	11.4	19.5	13.0
1970	11.5	11.5	19.4	13.1
1971	11.6	11.5	19.2	12.9
1972	11.8	11.6	18.7	12.5
1973	11.9	11.8	18.4	12.3
1974	12.1	11.9	18.5	12.2
1975	12.1	12.0	18.4	12.0
1976	12.2	12.0	18.1	11.8
1977	12.3	12.1	18.0	11.8
1978	12.4	12.2	17.8	11.6
1979	12.4	12.3	17.5	11.6
1980	12.5	12.4	17.6	11.7
1981	12.6	12.5	17.6	11.6
1982	12.8	12.6	17.7	11.7
1983	12.9	12.7	17.6	11.7
1984	12.9	12.8	17.6	11.7
1985	12.9	12.8	17.5	11.7
1986	12.9	12.9	17.6	11.7
1987	13.0	12.9	17.5	11.8
1988	13.0	12.9	17.6	12.0
1989	13.0	13.0	17.7	12.0
1990	13.0	13.0	17.8	12.1

[1] Averages are calculated by weighting educational attainment or work experience by the hours of workers.

So the 0.3-percentage-point growth rate of labor composition implies that changes in the education and experience of the work force contributed approximately 0.2 percentage point to the 2.3-percent annual growth rate of labor productivity. (See table 23 of chapter III.)[9] In other words, *changes in the skill*

[9] Labor and multifactor productivity measures presented in this bulletin are based on the August 1991 press release "Multifactor Productivity Measures, 1990" (USDL 91-412). Output and capital measures do not reflect the December 1991 comprehensive revisions to the National Income and Product Accounts. The output and capital investment data BLS used to prepare measures of capital input are therefore aggregated using 1982 price weights. The aggregates in the December 1991 revisions are prepared using 1987 price weights. Hours data include minor revisions subsequent to the August 1991 press release.

Table 2. Average annual rates of growth of labor input, hours of all persons, and labor composition, selected periods, 1948-90[1]
(Percent per year)

Period	Labor input	Hours of all persons	Labor composition
Private business			
1948-90	1.3	1.0	0.3
1948-73	.9	.6	.3
1973-90	1.9	1.6	.3
1973-79	1.7	1.7	.0
1979-90	2.0	1.5	.5
Private nonfarm business			
1948-90	1.7	1.4	0.3
1948-73	1.4	1.2	.2
1973-90	2.1	1.7	.3
1973-79	1.9	1.9	.0
1979-90	2.2	1.7	.5

Note: Growth rates of hours and labor composition may not sum to the growth rate of labor input due to rounding.

[1] Labor input and labor composition are Tornqvist indexes.

level of the work force account for about 9 percent of the growth in output per hour since 1948.

The growth of labor composition in private business has proceeded at an irregular pace over the last 40 years (table 2). Prior to 1973, labor composition grew 0.3 percent annually. From 1973 to 1979, labor composition remained unchanged. A slower growth rate for labor productivity was recorded for the 1973-79 period, but the slower labor composition growth rate can only explain a small fraction of the decline in this rate. (See table 23 in chapter III.) Between 1979 and 1990, labor composition grew 0.5 percent annually. Labor productivity continued to make small advances after 1979. Without the faster labor composition growth of this period, labor productivity would have posted even smaller gains than in the 1973-79 period.

The trends in labor input for private nonfarm business are very similar to those for private business (table 2). The Tornqvist index of labor input grew 1.7 percent annually between 1948 and 1990. The growth rate of labor composition was 0.3 percent, the same as in private business. Labor composition growth contributed 0.2 percentage point to the 1.9-percent annual growth rate of output per hour over the period. (See table 24 in chapter III).

The annual average growth rate of labor composition in private nonfarm business declined from 0.2 percent during the period 1948-73 to 0.0 percent between 1973 and 1979. Again, the declining growth rate of labor composition explains only a small fraction of the labor productivity slowdown. After 1979, labor composition grew at its fastest rate of the postwar period, 0.5 percent. Labor productivity would have increased more slowly after 1979 compared to the 1973-79 period if not for the rapid gains in labor composition.

Organization of the bulletin

The remainder of this bulletin is divided into chapters II and III and appendixes, A through H. Chapter II presents the labor composition measures developed for this study. Changes in the educational attainment, age distribution, and level of work experience in the work force are examined. The effects of these shifts are incorporated into new measures of labor input and labor composition. Chapter III examines labor and multifactor productivity growth in the U.S. private business and private nonfarm business sectors from 1948 to 1990 and the factors contributing to it, especially labor composition growth.

Appendix A presents the basic economic model for measuring productivity and the contribution of labor composition to productivity growth. Choosing the appropriate set of worker characteristics for measuring labor composition is discussed. These traits are then related to earnings to yield a pricing function for labor services (wage equation). The growth accounting model used for productivity analysis is expanded to include changes in labor composition. Appendix B describes some of the empirical tests of wage equations that were performed using the National Longitudinal Survey (NLS). These tests were conducted to determine the effects of certain work force characteristics on estimated wage equations. Appendix C discusses the derivation and estimation of the experience equation developed for this study. Appendix D describes the derivation of the annual hours matrices which are cross-classified by education and experience for each sex. Appendix E describes in detail the estimation and derivation of the earnings function used to derive weights for each type of worker hours. Estimated parameters for the effects of education and experience on earnings are presented, and the relative earnings of workers are examined. Appendix F analyzes the sensitivity of the measures of labor composition to a number of the assumptions and choices made in specifying the labor composition model. Appendix G compares the measures presented in this study with measures produced by other researchers. Finally, appendix H attempts to separate the effects of education from work experience. Conditions for an exact decomposition are presented. An exact decomposition is highly improbable, and the separate effects presented in this appendix must be regarded as estimates.

Chapter II. Labor Input and Labor Composition Growth, 1948-90

This chapter presents measures of labor input allowing for a world populated with workers of varying skills. The chapter begins by examining why workers differ in their level of skill. A method of measuring the flow of labor services provided by each kind of worker is then developed. In addition, two important differences among workers, education and work experience, are explored further. Trends in the work force are reviewed including educational attainment and work experience.

This chapter goes on to describe a model of labor input measurement with diverse or heterogeneous workers. This model is based on the neoclassical theory of the firm and relies on human capital theory to identify sources of worker heterogeneity. Labor input growth is then divided between the growth in total labor hours and labor composition. Labor composition gauges how changes in the educational attainment and work experience of the labor force affect measures of labor input. A discussion of the concept of labor composition follows. Next, the BLS methodology is briefly discussed; a complete description appears in appendixes A, C, D, and E. Finally, labor input and labor composition measures for the period 1948-90 are presented.

Chapter III places labor composition within a productivity framework and examines how changes in labor composition have affected productivity and economic growth.

Heterogeneous labor

Labor productivity measures have traditionally defined labor input as the sum of all hours worked by employees, proprietors, and unpaid workers. As a result, an hour worked by a highly experienced surgeon and an hour worked by a newly hired teenager at a fast food restaurant are treated as equal amounts of labor. It does not matter who was actually working or what kinds of jobs workers held. All workers are treated as if they are identical.

In this study, workers are treated as heterogeneous.[1] Differences in skills imply that different types of workers provide different amounts of labor services. Instead of adding together the hours of all workers, the hours of each type of worker are considered a separate input. Aggregate labor input is measured by combining the flows of services provided by each type of labor. The flow of services for any type of labor, in turn, depends on the number of hours and the level of skills of the workers.

The concept of labor services comes from the theory of the firm. Labor services are measured by the value of the marginal product of labor. The value of the marginal product of each type of labor is defined as the extra value of output produced by an additional unit (hour) of that type of labor holding all other inputs constant. That is, the value of the marginal product for each type of worker measures the value produced in an additional hour of work.

In the real world, it is extremely difficult if not impossible to observe directly the marginal products of individual workers. In a perfectly competitive economy, however, inputs are paid the values of their marginal products. Therefore, the relative service flows can be estimated from the hourly wage rates of each kind of worker.

Because it is impractical to distinguish between workers based on every observable earnings difference, human capital theory is used as an organizing principle to help distinguish between labor inputs. Articles in the human capital literature as well as the hedonic wage literature have consistently shown that earnings differ by education and work experience. Hence, wage differentials due to education and work experience, for each sex, are used in the construction of labor input measures.

Human capital theory. Human capital theory explains differences in the level of skills across the work force. This theory formalizes the intuitive notion that workers improve or learn new skills through education and job training.[2] These learning activities are an investment because workers sacrifice current earnings for higher wages in the future. The set of skills acquired through education and training makes up the worker's stock of human capital.

Under most circumstances, firms will also invest to train their employees. Most company training is informal, however, and most companies do not prepare data on the training investments they make in their workers. Consequently, practical considerations prevent one from distinguishing between workers based on training investments. Related research has demonstrated that the amount of training can relate to the lev-

[1] Previous studies of this issue include D. Jorgenson, F. Gollop, and B. Fraumeni, *Productivity and U.S. Economic Growth*, Harvard University Press (Cambridge, MA., 1987); Edward Denison, *Trends in American Economic Growth, 1929-82*, Brookings Institution (Washington, DC, 1985); and Peter Chinloy, "Sources of Quality Change in Labor Input," *American Economic Review*, March 1980, pp. 108-119.

[2] Gary Becker, *Human Capital*, Columbia University Press, (New York 1975).

el of a worker's job experience.[3] In this study, hours of labor are classified by education and work experience for each sex to measure the flow of labor services. For a more thorough discussion of the human capital model and its usefulness in measuring labor input, see appendix A.

Trends in employment and average weekly hours

Change in labor input depends on the change in employment, change in the length of the workweek, and change in the education and experience distributions of the work force. The Bureau's measures of total hours incorporate the results of the Hours At Work survey. As a result, the concept of total hours has been converted from an hours paid basis to hours at work. The hours of paid leave are now excluded from the measures of labor.

Chart 1 shows that the hours at work of all persons in private business have grown at an uneven pace over the last 42 years. Besides the obvious fluctuations in total hours with the business cycle, after 1973 hours grew more than twice as fast as in the earlier period. Prior to 1973, hours of all persons increased at an average annual rate of 0.6 percent but grew 1.6 percent annually thereafter. The faster growth rate in total hours in the later period is almost completely due to an increase in the growth rate in employment from 1.1 percent to 2.0 percent. The average workweek declined at about the same rate in both periods: 0.5 percent in 1948-73 and 0.4 percent in 1973-90.

Total hours at work in the U.S. private business sector grew by 51 percent between 1948 and 1990. As indicated in table 3, the sources of this growth come from two opposing trends. Employment increased nearly 83 percent over the period, and the average workweek declined from 40.9 to 33.9 hours or 17 percent.[4]

Employment growth has been concentrated primarily in the nongoods-producing industries. This sector, which includes services, transportation, communications, utilities, wholesale and retail trade, finance, insurance, and real estate, accounted for 97 percent of the increase in total employment (table 4). The growth of nongoods-producing employees has been the predominant source of employment growth since 1948. Employees in this sector accounted for 91 percent of total employment growth, while proprietors and unpaid family workers accounted for only 6 percent. The goods-producing industries of agriculture, mining, construction, and manufacturing have contributed only 3 percent of the increase in total employment since 1948 and essentially none after 1973 (table 5). The number of employees in the goods-producing industries increased modestly between 1948 and 1990 and accounted for 13 percent of the increase in total employment. In contrast, the number of proprietors and unpaid family workers in goods-producing industries declined over the period to less than half the 1948 level.

[3] Jacob Mincer, *Schooling, Experience and Earnings*, Columbia University Press, (New York, 1974).

[4] The average workweek is an average over full-time and part-time workers and is measured in hours at work rather than hours paid.

Table 3. Hours, employment, and average weekly hours in private business, 1948-90[1]

Year	Total hours	Employment	Average weekly hours
1948	109.7	51.5	40.9
1949	105.9	50.3	40.5
1950	106.9	50.8	40.5
1951	109.9	52.3	40.4
1952	109.8	52.5	40.2
1953	110.8	53.4	39.9
1954	107.0	52.0	39.6
1955	110.9	53.6	39.8
1956	112.6	54.8	39.5
1957	110.8	54.7	39.0
1958	105.5	52.5	38.7
1959	109.6	54.1	38.9
1960	109.6	54.4	38.8
1961	107.8	53.8	38.6
1962	109.5	54.4	38.7
1963	110.0	54.8	38.6
1964	111.7	55.8	38.5
1965	115.2	57.4	38.6
1966	117.9	59.1	38.4
1967	117.5	59.9	37.7
1968	119.1	61.1	37.5
1969	121.9	63.1	37.2
1970	119.5	62.9	36.6
1971	118.9	62.8	36.4
1972	122.9	65.0	36.4
1973	127.3	67.7	36.2
1974	127.3	68.7	35.6
1975	121.9	66.6	35.2
1976	125.4	68.6	35.1
1977	130.3	71.5	35.0
1978	136.7	75.4	34.9
1979	141.0	78.2	34.7
1980	139.6	78.4	34.3
1981	140.6	79.2	34.1
1982	137.0	77.8	33.9
1983	139.5	78.6	34.1
1984	147.5	82.6	34.3
1985	150.6	84.8	34.2
1986	151.6	86.2	33.8
1987	156.3	88.6	33.9
1988	161.5	91.3	34.0
1989	165.7	93.4	34.1
1990	165.6	94.0	33.9

[1] Excludes government enterprises.

NOTE: Total hours are expressed in billions of hours worked. Employment is expressed in millions of persons. Average weekly hours are expressed in hours worked per worker per week.

The average workweek in both sectors decreased over the period. In the nongoods-producing sector, the workweek fell sharply from 40.3 hours in 1948 to 32.3 hours in 1990. The workweek in the goods-producing sector declined moderately from 41.4 hours in 1948 to 37.9 hours in 1990. Even this small decrease in the workweek of the goods-producing sector was enough to reduce total hours between 1948 and 1990. Consequently, the growth of hours at work in private business has been entirely confined to the nongoods-producing sector (chart 2).

Chart 1. Indexes of total hours, employment, and average weekly hours in private business, 1948-90

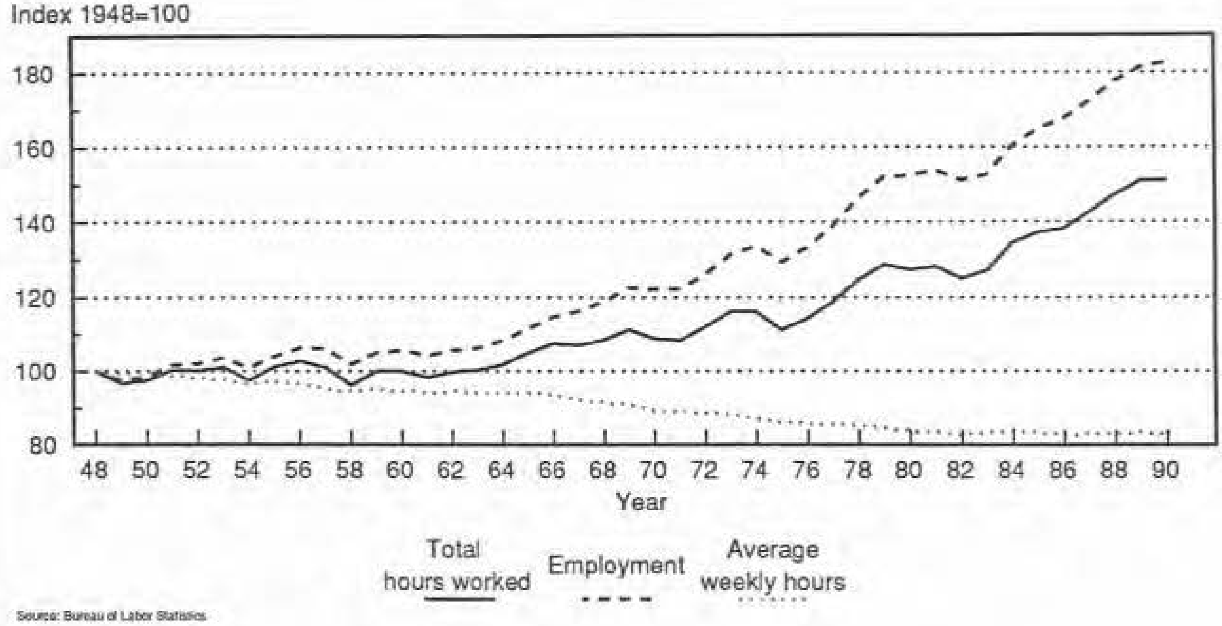


Chart 2. Sources of hours growth in private business, goods- and nongoods-producing sectors, 1948-90

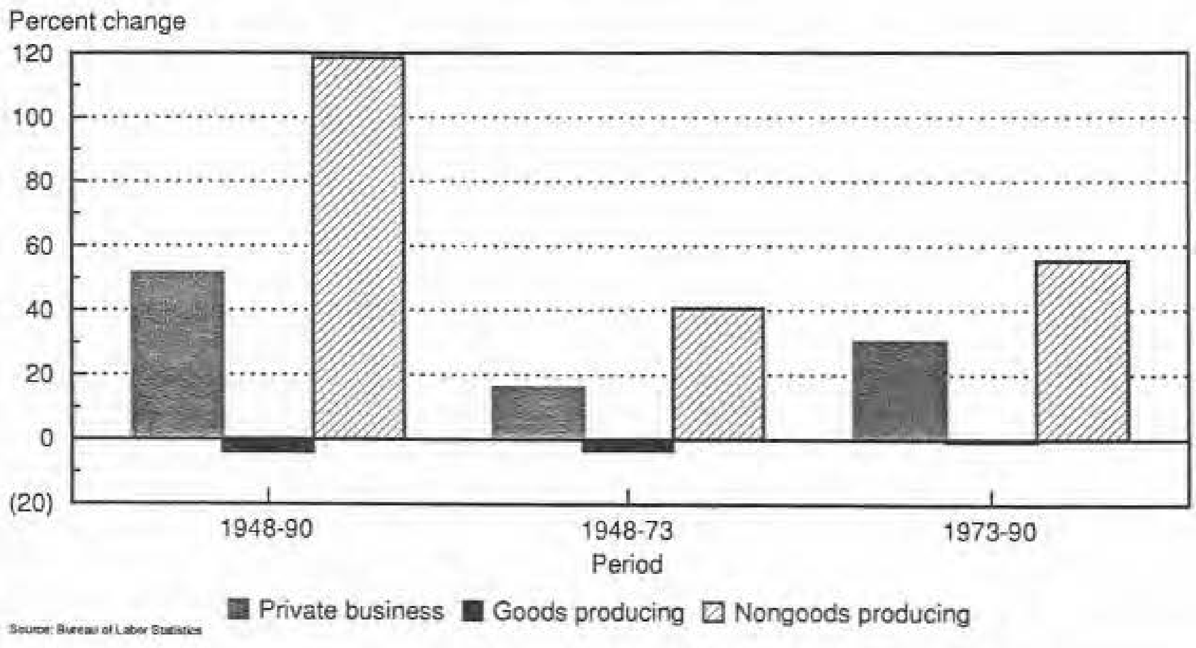


Table 4. Hours, employment, and average weekly hours in nongoods-producing industries, 1948-90

Year	Total[1] hours	Employment[1]			Average weekly hours
		All workers	Employees only	Other[2] employment	
1948	51.55	24.57	19.74	4.83	40.3
1949	51.11	24.48	19.61	4.87	40.2
1950	51.13	24.63	19.86	4.77	39.9
1951	52.24	25.24	20.63	4.61	39.8
1952	52.66	25.64	21.17	4.47	39.5
1953	53.35	26.12	21.56	4.56	39.3
1954	53.33	26.13	21.48	4.65	39.2
1955	54.84	26.82	22.09	4.73	39.3
1956	56.16	27.58	22.76	4.82	39.2
1957	56.34	27.95	22.96	4.99	38.8
1958	55.40	27.59	22.56	5.03	38.6
1959	57.01	28.35	23.24	5.11	38.7
1960	57.91	28.84	23.64	5.20	38.6
1961	57.83	29.02	23.74	5.28	38.3
1962	58.46	29.39	24.25	5.14	38.3
1963	59.03	29.79	24.75	5.04	38.1
1964	60.45	30.63	25.49	5.14	38.0
1965	62.01	31.55	26.44	5.11	37.8
1966	63.14	32.51	27.49	5.02	37.3
1967	63.60	33.30	28.45	4.85	36.7
1968	64.84	34.23	29.46	4.77	36.4
1969	67.07	35.72	30.79	4.93	36.1
1970	67.86	36.58	31.70	4.88	35.7
1971	68.64	37.30	32.28	5.02	35.4
1972	70.42	38.53	33.47	5.06	35.1
1973	72.55	39.99	34.90	5.09	34.9
1974	73.53	41.09	35.89	5.20	34.4
1975	73.32	41.36	36.10	5.26	34.1
1976	75.02	42.68	37.45	5.23	33.8
1977	77.80	44.60	39.10	5.50	33.5
1978	81.39	47.02	41.37	5.65	33.3
1979	84.36	49.08	43.16	5.92	33.1
1980	85.29	50.05	44.02	6.03	32.8
1981	86.56	51.07	44.94	6.13	32.6
1982	87.27	51.45	45.06	6.39	32.6
1983	89.58	52.66	46.04	6.62	32.7
1984	94.32	55.38	48.64	6.74	32.8
1985	97.60	57.63	50.94	6.69	32.6
1986	99.58	59.46	52.82	6.64	32.2
1987	103.96	61.74	54.73	7.01	32.4
1988	107.09	64.04	56.80	7.24	32.2
1989	110.86	65.92	58.56	7.36	32.3
1990	112.75	67.13	59.88	7.45	32.3

[1] Includes government enterprises.
[2] Proprietors and unpaid family workers.
NOTE: Total hours are expressed in billions of hours worked. Employment is expressed in millions of persons. Average weekly hours are expressed in hours per worker per week.

Table 5. Hours, employment, and average weekly hours in goods-producing industries, 1948-90

Year	Total[1] hours	Employment[1]			Average weekly hours
		All workers	Employees only	Other[2] employment	
1948	60.07	27.87	20.43	7.44	41.4
1949	56.86	26.76	19.33	7.43	40.9
1950	57.89	27.18	20.17	7.01	41.0
1951	59.86	28.10	21.51	6.59	41.0
1952	59.45	28.01	21.61	6.40	40.8
1953	59.69	28.33	22.34	5.99	40.5
1954	55.99	26.95	21.00	5.95	40.0
1955	58.52	27.92	22.00	5.92	40.3
1956	58.85	28.37	22.57	5.80	39.9
1957	56.90	27.90	22.42	5.48	39.2
1958	52.61	26.07	20.93	5.14	38.8
1959	55.11	26.96	21.84	5.12	39.3
1960	54.33	26.80	21.97	4.83	39.0
1961	52.62	26.01	21.30	4.71	38.9
1962	53.71	26.35	21.81	4.54	39.2
1963	53.74	26.30	22.02	4.28	39.3
1964	54.05	26.48	22.31	4.17	39.3
1965	56.10	27.23	23.16	4.07	39.6
1966	57.60	28.00	24.24	3.76	39.6
1967	56.80	28.04	24.45	3.59	39.0
1968	57.24	28.38	24.81	3.57	38.8
1969	57.93	28.86	25.37	3.49	38.6
1970	54.89	27.88	24.53	3.35	37.9
1971	53.58	27.16	23.89	3.27	37.9
1972	55.66	28.01	24.70	3.31	38.2
1973	58.02	29.28	25.95	3.33	38.1
1974	57.16	29.28	25.94	3.34	37.5
1975	52.04	26.93	23.69	3.24	37.2
1976	53.84	27.64	24.45	3.19	37.5
1977	56.00	28.66	25.44	3.22	37.6
1978	58.89	30.11	26.74	3.37	37.6
1979	60.32	30.94	27.53	3.41	37.5
1980	58.10	30.16	26.69	3.47	37.0
1981	57.74	29.97	26.56	3.41	37.0
1982	53.48	28.20	24.88	3.32	36.5
1983	53.74	27.81	24.54	3.27	37.2
1984	57.17	29.15	25.89	3.26	37.7
1985	57.14	29.15	25.97	3.18	37.7
1986	56.37	28.85	25.62	3.23	37.6
1987	56.70	28.96	25.81	3.16	37.7
1988	58.32	29.44	26.23	3.21	38.1
1989	58.59	29.55	26.37	3.20	38.1
1990	57.67	29.20	26.01	3.19	37.9

[1] Includes government enterprises.
[2] Proprietors and unpaid family workers.
NOTE: Total hours are expressed in billions of hours worked. Employment is expressed in millions of persons. Average weekly hours are expressed in hours per worker per week.

Trends in educational attainment

The educational attainment of men and women in the work force has increased greatly over the last 42 years. Tables 6 and 7 indicate that the proportion of hours worked by high school graduates rose substantially over the period. (In this study, the measures of educational attainment are based on the distribution of hours rather than the distribution of employment or the civilian labor force.) In addition, the portion of hours worked by men with at least a college degree jumped from 6.0 to 25.2 percent of the male total over the period. Women with at least a college degree showed equally spectacular gains with the proportion rising from less than 4 percent to

Table 6. Distribution of all hours worked by men in the private business sector by years of completed schooling, 1948-90
(in percent)

Year	Years of completed schooling						
	0-4	5-8	9-11	12	13-15	16	17+
1948	8.3	35.6	20.5	23.1	6.5	3.6	2.4
1949	9.3	36.0	19.5	21.4	7.1	4.0	2.7
1950	9.1	35.9	19.5	21.5	7.2	4.1	2.7
1951	8.8	35.1	19.1	22.4	7.4	4.3	2.9
1952	8.5	34.4	18.8	23.3	7.6	4.4	3.0
1953	8.1	33.4	19.0	24.0	7.7	4.7	3.2
1954	7.6	32.4	19.1	24.7	7.7	4.9	3.4
1955	7.1	31.3	19.4	25.7	7.8	5.1	3.6
1956	6.7	30.3	19.6	26.5	7.9	5.3	3.7
1957	6.2	29.3	19.8	27.3	7.9	5.5	3.9
1958	5.9	28.9	20.2	26.8	8.6	5.5	4.0
1959	5.5	28.4	20.7	26.4	9.3	5.5	4.0
1960	5.2	27.8	21.1	26.1	10.1	5.6	4.1
1961	4.8	25.8	20.4	28.2	10.2	6.5	4.2
1962	4.5	23.8	19.6	30.2	10.2	7.3	4.3
1963	4.1	22.9	19.6	31.5	10.2	7.2	4.5
1964	3.7	22.0	19.6	32.8	10.2	7.0	4.7
1965	3.8	21.1	19.3	33.8	10.1	7.5	4.5
1966	3.4	20.4	19.5	34.4	10.2	7.8	4.4
1967	2.9	18.6	18.8	35.3	11.9	7.6	5.1
1968	2.7	17.9	18.7	35.8	12.2	7.7	5.0
1969	2.5	17.0	17.7	36.4	12.8	8.2	5.4
1970	2.4	15.7	16.9	37.2	13.5	8.6	5.7
1971	2.3	14.8	17.2	37.2	13.8	8.8	5.8
1972	2.3	12.9	16.3	38.8	14.6	9.1	6.0
1973	2.1	12.4	15.7	38.6	15.2	9.5	6.5
1974	1.7	10.7	14.9	38.7	16.0	10.9	7.2
1975	1.8	10.8	14.9	38.9	15.9	10.7	7.0
1976	1.7	10.1	15.0	38.5	16.4	10.9	7.3
1977	1.7	9.7	14.5	38.4	17.4	10.9	7.5
1978	1.5	8.9	13.6	39.0	18.1	11.1	7.8
1979	1.4	8.5	13.7	39.0	17.8	11.5	8.1
1980	1.4	7.8	13.1	39.5	17.8	12.1	8.3
1981	1.3	7.3	12.5	39.6	17.6	12.5	9.2
1982	1.3	6.5	11.7	38.6	18.0	13.8	10.2
1983	1.0	6.4	10.9	39.2	18.4	13.9	10.1
1984	1.1	6.2	11.0	39.2	18.7	14.1	9.8
1985	1.1	5.9	10.5	39.3	19.2	14.5	9.5
1986	1.0	5.6	10.7	38.9	19.3	14.5	10.0
1987	1.1	5.1	10.6	39.3	18.8	14.8	10.2
1988	1.2	5.1	10.2	38.8	19.5	14.8	10.5
1989	1.3	4.9	10.0	38.7	19.9	15.1	10.1
1990	1.3	4.5	9.7	39.1	20.2	15.1	10.1

NOTE: Total may not sum to 100.0 due to rounding.

Table 7. Distribution of all hours worked by women in the private business sector by years of completed schooling, 1948-90
(in percent)

Year	Years of completed schooling						
	0-4	5-8	9-11	12	13-15	16	17+
1948	4.8	29.5	18.8	36.3	6.8	2.5	1.3
1949	4.5	25.4	20.5	35.5	9.2	3.2	1.6
1950	4.5	25.4	20.3	35.7	9.2	3.2	1.6
1951	4.6	25.8	20.1	36.4	8.4	3.2	1.6
1952	4.6	26.2	19.8	37.0	7.6	3.2	1.6
1953	4.3	25.5	19.9	37.9	7.7	3.2	1.6
1954	4.1	24.7	19.8	38.8	7.8	3.3	1.6
1955	3.7	24.1	19.8	39.6	7.8	3.3	1.6
1956	3.4	23.4	19.9	40.4	7.9	3.4	1.6
1957	3.1	22.5	19.8	41.3	8.0	3.5	1.7
1958	3.0	22.3	20.6	40.2	8.8	3.4	1.6
1959	2.9	22.5	21.7	38.7	9.5	3.2	1.5
1960	2.7	21.7	22.2	38.0	10.6	3.3	1.5
1961	2.5	20.5	20.8	40.3	10.5	4.1	1.2
1962	2.3	18.9	19.4	43.0	10.6	4.8	1.0
1963	2.2	18.5	19.4	44.2	10.1	4.4	1.2
1964	2.1	18.2	19.4	45.2	9.7	4.0	1.4
1965	2.0	17.4	19.2	45.9	9.7	4.2	1.6
1966	1.6	16.7	19.1	47.4	10.3	3.5	1.5
1967	1.3	15.1	18.9	48.0	12.0	3.5	1.1
1968	1.4	13.5	18.6	50.2	11.7	3.3	1.3
1969	1.1	12.4	17.7	50.6	12.3	4.1	1.7
1970	1.1	11.7	17.2	50.3	13.3	4.3	2.1
1971	1.2	10.5	16.9	51.4	13.6	4.8	1.5
1972	1.1	9.5	16.0	52.5	14.2	4.8	1.8
1973	1.0	8.9	15.5	50.8	15.9	5.6	2.2
1974	.9	8.3	14.9	50.8	16.2	6.3	2.8
1975	.8	7.7	15.1	50.3	16.4	6.9	2.8
1976	.9	7.3	14.9	50.4	16.8	7.1	2.7
1977	.8	6.9	14.4	50.2	17.6	7.3	2.8
1978	.8	6.0	13.1	50.4	18.5	7.7	3.4
1979	.7	5.7	12.4	50.2	18.7	8.6	3.6
1980	.6	5.2	11.9	50.0	19.8	8.9	3.6
1981	.7	4.8	11.3	49.8	19.7	9.3	4.4
1982	.6	4.4	10.4	48.7	20.6	10.4	4.9
1983	.6	3.9	9.9	48.4	21.0	11.0	5.2
1984	.6	3.7	9.7	47.3	22.0	11.6	5.1
1985	.5	3.4	9.1	47.0	22.6	12.2	5.2
1986	.4	3.4	8.9	46.7	22.7	12.3	5.5
1987	.7	3.1	9.1	45.8	23.1	12.6	5.6
1988	.7	2.8	9.0	45.7	22.7	13.2	5.9
1989	.6	2.8	8.6	45.0	23.3	13.5	6.3
1990	.7	2.7	8.4	44.3	23.5	14.1	6.3

NOTE: Total may not sum to 100.0 due to rounding.

more than 20 percent.

The educational attainment of men and women differs in one important respect. Women in the work force are more likely to complete high school but less likely to continue on to college. In 1948, men without a high school diploma worked 64.4 percent of men's hours, but women without a high school diploma accounted for 53.1 percent of women's hours. Men with at least some college provided 12.5 percent of all men's hours, and comparable women supplied just 10.6 percent of women's hours. Even after 42 years of rising educational attainment, the same pattern remains. For men, 15.5 percent of hours were worked by those without a high school diploma, but for women the corresponding percentage is only 11.8. Men with at least some college supplied 45.4 percent of men's hours, and women with at least some college provided 43.9 percent of women's hours.

Table 8 indicates that the average educational attainment of women in the work force has increased from less than 10 years in 1948 to 13 years of completed schooling in 1990. The educational attainment of men has risen even more rapidly over the same period increasing from 9.3 years to 13.0 years of completed schooling. The educational attainment of all workers appears similar to the trend for men because men worked most of the hours in private business, especially early in the

Table 8. Average years of completed schooling of working men and women in the private business sector[1], 1948-90

Year	Men	Women	All workers
1948	9.3	9.9	9.4
1949	9.2	10.2	9.4
1950	9.2	10.2	9.4
1951	9.3	10.1	9.5
1952	9.4	10.1	9.6
1953	9.5	10.2	9.7
1954	9.7	10.2	9.8
1955	9.8	10.3	9.9
1956	9.9	10.4	10.0
1957	10.0	10.5	10.1
1958	10.1	10.5	10.2
1959	10.3	10.6	10.4
1960	10.2	10.5	10.3
1961	10.4	10.7	10.5
1962	10.6	10.8	10.7
1963	10.7	10.8	10.8
1964	10.8	10.9	10.8
1965	10.9	10.9	10.9
1966	10.9	11.0	11.0
1967	11.0	11.0	11.0
1968	11.2	11.2	11.2
1969	11.4	11.4	11.4
1970	11.5	11.5	11.5
1971	11.6	11.5	11.6
1972	11.8	11.6	11.7
1973	11.9	11.8	11.8
1974	12.1	11.9	12.1
1975	12.1	12.0	12.1
1976	12.2	12.0	12.1
1977	12.3	12.1	12.2
1978	12.4	12.2	12.3
1979	12.4	12.3	12.4
1980	12.5	12.4	12.5
1981	12.6	12.5	12.6
1982	12.8	12.6	12.7
1983	12.9	12.7	12.8
1984	12.9	12.8	12.8
1985	12.9	12.8	12.9
1986	12.9	12.9	12.9
1987	13.0	12.9	12.9
1988	13.0	12.9	13.0
1989	13.0	13.0	13.0
1990	13.0	13.0	13.0

[1] Averages are calculated by weighting educational attainment by hours of workers.

Table 9. Mean age of civilian labor force aged 16 or older, by sex, 1948-90

Year	Mean age in years		
	Women	Men	All workers
1948	36.6	40.0	39.0
1950	37.5	40.1	39.4
1955	39.0	41.1	40.4
1958	39.6	40.9	40.5
1960	39.8	40.7	40.4
1965	39.5	40.3	40.0
1970	38.4	39.7	39.2
1975	36.8	38.1	37.6
1980	36.3	37.4	36.9
1985	36.5	37.5	37.0
1990	37.4	37.9	37.7

Source: Calculated from data in *Employment and Earnings*, published by the U.S. Department of Labor, Bureau of Labor Statistics, January 1991 and earlier issues.

period. For both men and women, the increase in educational attainment has been fairly steady over most of the period.

Trends in the age and work experience of the work force

The declining level of estimated work experience is a countervailing trend which has moderated the increase in the average level of skills and labor input. An important factor influencing the level of estimated work experience is the age distribution of the work force. Table 9 shows a drop in the average age of the civilian labor force of more than 1 year since 1948. The average age of the civilian labor force increased from 1948 to 1958 but then declined over a prolonged period which lasted about two decades. Around 1980, the average age began to increase slightly once again. This general pattern of movements in the average age is evident for both men and women. A salient difference between the two series is that the mean age of women in the labor force was higher in 1990 than in 1948, but the mean age of men in the labor force was considerably lower in 1990 than in 1948.

The primary force behind the decline in the average age of the labor force was the entrance into the labor market of members of the baby boom generation (table 10). Prior to the entrance of the baby boom generation, the baby bust of the depression era was the primary influence on the age distribution of the labor force. Between 1948 and 1960, the proportions of the work force age 16-24 and age 25-34 decreased, while the proportion of the work force over age 35 steadily increased. From 1960 to 1975, the percentage of the labor force in the 16-24 age group rose from 16.6 to 24.1. Table 10 also shows that the aging of the baby boomers is the primary force behind the rise in the average age of the American worker in the 1980's. From 1980 to 1990, the proportion of the labor force between the ages of 35 and 44 jumped from 19.1 to 25.2 percent.

The average years of estimated work experience shown in table 11 exhibit a pattern similar to the trends in the average

age of the labor force.[5] (Again, the averages are based on the distribution of hours and not employment.) The average number of years of estimated work experience increased from 17.4 in 1948 to 18.4 in 1958. The average declined to 15.5 years in 1979; however, unlike the average age, the average experience of all workers did not rise between 1980 and 1990.

Women have, on average, less work experience than men. In 1948, men had 6.4 years more work experience than women, but this gap had narrowed to 5.7 years by 1990. For both men and women, the average years of work experience increased in the 1950's and peaked in 1958. The men's average fell gradually in the 1960's and more rapidly in the 1970's; since 1979, it has varied little. The average years of work experience for women also dropped in the 1960's and 1970's, but it rose between 1980 and 1990, from 11.7 to 12.1 years. The net effect of all these changes was that average work experience was 1 year lower for men and negligibly lower for women in 1990 than in 1948.

The next section describes a method for measuring changes in the average skill level or composition of the work force. Labor composition accounts not only for changes in the educational attainment and experience of the work force but also for changes in the relative value, based on prices paid, of the underlying skills.

Table 10. Percent distribution of civilian labor force aged 16 or older, 1948-90

Year	Age group				
	16-24	25-34	35-44	45-54	55+
1948	19.51	23.52	22.10	18.00	16.87
1950	18.52	23.50	22.43	18.40	17.15
1955	15.01	23.16	23.68	19.98	18.18
1958	15.57	21.69	23.70	20.96	18.08
1960	16.58	20.66	23.37	21.33	18.07
1965	19.03	19.12	22.62	21.16	18.08
1970	21.56	20.58	19.86	20.48	17.52
1975	24.12	24.38	18.03	18.22	15.25
1980	23.66	27.33	19.14	15.81	14.06
1985	20.46	29.06	22.58	15.00	12.90
1986	19.83	29.36	23.11	15.05	12.65
1987	19.16	29.39	23.74	15.19	12.51
1988	18.52	29.18	24.19	15.70	12.40
1989	17.87	28.98	24.70	16.08	12.37
1990	17.03	28.69	25.52	16.42	12.34

Source: U.S. Department of Labor, Bureau of Labor Statistics, *Employment and Earnings*, January 1991 and earlier issues.

[5] The measures of work experience are derived from a function relating work histories reported to the Social Security Administration to the individual's age, education, marital status, children, and other traits. To the extent that changes in labor force behavior can be modelled by changes in the underlying characteristics of the work force, the function accurately estimates work experience throughout the entire period. However, the work histories cover the period 1937-73, and changes by cohort in labor force participation, especially for women after 1973, may not be reflected in the work experience measures. See the section later in this chapter and appendix C for a description of the work experience estimates.

Table 11. Average number of years of estimated work experience for all workers, and for men and women, in private business, 1948-90

Year	All workers	Men	Women
1948	17.4	18.8	12.4
1949	17.5	19.0	12.5
1950	17.7	19.1	12.5
1951	17.8	19.3	12.7
1952	17.9	19.4	12.9
1953	18.1	19.6	13.0
1954	18.3	19.8	13.1
1955	18.2	19.7	13.3
1956	18.2	19.8	13.5
1957	18.3	19.8	13.6
1958	18.4	19.9	13.7
1959	18.3	19.8	13.7
1960	18.2	19.8	13.4
1961	18.2	19.8	13.6
1962	18.2	19.8	13.5
1963	18.2	19.8	13.6
1964	18.2	19.7	13.6
1965	18.1	19.6	13.5
1966	18.0	19.6	13.4
1967	17.9	19.6	13.2
1968	17.8	19.6	13.1
1969	17.7	19.5	13.0
1970	17.6	19.4	13.1
1971	17.4	19.2	12.9
1972	17.0	18.7	12.5
1973	16.6	18.4	12.3
1974	16.6	18.5	12.2
1975	16.4	18.4	12.0
1976	16.1	18.1	11.8
1977	16.0	18.0	11.8
1978	15.8	17.8	11.6
1979	15.5	17.5	11.6
1980	15.5	17.6	11.7
1981	15.5	17.6	11.6
1982	15.5	17.7	11.7
1983	15.5	17.6	11.7
1984	15.4	17.6	11.7
1985	15.4	17.5	11.7
1986	15.4	17.6	11.7
1987	15.3	17.5	11.8
1988	15.4	17.6	12.0
1989	15.5	17.7	12.0
1990	15.6	17.8	12.1

Labor input measurement

This study introduces a more refined measure of labor input. Previously, the multifactor productivity measures for private business and private nonfarm business published by the Bureau of Labor Statistics measured labor input as the sum of total hours of all workers. Changes in labor composition are not captured by a measure of total hours because such a measure does not account for differences among workers. Instead, changes in labor composition show up in the measures of productivity change.

In this bulletin, the differences among workers are explicitly recognized and taken into account. Equation 1 is a shorthand description of how the hours of different types of workers are aggregated into a single measure of labor input.

(1) $L = g(h_1, h_2, \ldots, h_n)$

In equation 1, the aggregate measure of labor input, L, combines the hours of each of the n types of labor, h_1 through h_n.[6] Under certain assumptions regarding the function in equation 1, the growth rate of labor input can be measured using a Tornqvist index.[7] The Tornqvist index aggregates the changes in the hours of different types of workers, in the following manner:

$$(2)\quad \ln(L_t / L_{t-1}) = \Sigma_i\, [(s_{i,t} + s_{i,t-1})/2] \ln(h_{i,t} / h_{i,t-1}),$$
$$= \Sigma_i\, w_{i,t} \ln(h_{i,t} / h_{i,t-1}),$$

where ln is the natural logarithm, t is a given year, t-1 is the previous year, and $s_{i,t}$ is the percentage change in L induced by a 1-percent change in $h_{i,t}$. As explained in appendix A, s_i equals the fraction (or share) of labor compensation that is paid to group i, under conditions of competitive markets. The weight $w_{i,t}$ is the average of $s_{i,t}$ and $s_{i,t-1}$; the weights sum to one.

Computation of labor composition growth

Labor composition growth shows the contribution to labor input of changes in the education and experience of the work force. The growth rate of labor composition, C, is defined as the difference between the growth rate of the Tornqvist index of labor input, L, and the growth rate of total hours, H. Using a lowercase boldface letter, x, for the growth rate of any variable X, the growth rate of labor composition can be expressed as:[8]

(3) $\mathbf{c} = \mathbf{l} - \mathbf{h}$.

The growth rate of labor composition, **c**, is thus the difference between the growth rate of labor input, **l**, and the growth rate of total hours (**h**).

Rearranging terms in equation 3 yields an expression for the growth rate of labor input:

(4) $\mathbf{l} = \mathbf{h} + \mathbf{c}$.

This equation points out that the growth of labor input is the combination of the growth of unadjusted hours and the growth of the composition of the work force. This implies that a 1-percent increase in labor composition has the same effect on the growth rate of labor input as a 1-percent increase in total hours.

Increases in labor composition result when labor input grows faster than total hours. The quantity of labor services per hour is increasing whenever $\mathbf{l} - \mathbf{h} > 0$. Increase in the quantity of services per hour is equivalent to an increase in the marginal product of labor of the average worker provided that the level of technology, all other inputs, and the quality of education and training are held fixed. More intuitively, the marginal product of labor of the average worker generally rises when the distribution of the work force changes so as to increase the average level of education or work experience.

As in the example in chapter 1, imagine replacing some of the inexperienced teenagers with an equal number of experienced surgeons. Experienced doctors have a higher marginal product than inexperienced teenagers (holding all else constant), therefore, the average skill level and marginal product of the work force will rise. The measures of labor composition reflect this rise because increases in the hours of higher paid workers such as doctors are weighted more heavily than lower paid workers. Weighted labor input increases while total hours of the work force remain unchanged, and so labor composition will increase.

For the measures of labor composition in this study, two sources of workers' skills are considered: education and work experience. Workers' skills may well vary with other characteristics, such as occupation and industry, but they are not included here. For brevity, the term "skills" will refer here to the combined effect on a worker's abilities of education and work experience.

Changes in wages can have an impact on labor composition growth through their effect on labor compensation shares. Recall that labor composition growth is a weighted average of the growth rates of the hours of each type of worker. Increasing the weight on rapidly growing hours or reducing the weight on slowly growing hours leads to faster labor input and labor composition growth. The weights can change from year to year because of shifts in the relative compensation of groups of workers. For example, the earnings of college graduates has increased faster than the earnings of high school graduates since the early 1970's. As a result, the share of compensation and the weight on the rapidly growing hours of college graduates has increased and spurred labor composition growth in the 1980's. However, if all wages increase at the same rate, these increases will have no impact on labor composition growth because the labor compensation shares or weights are not altered if everyone's wages change proportionately.

Development of weights and hours. The calculation of the Tornqvist indexes of labor input and composition requires information on the compensation and hours of workers. Estimated wages and hours are combined to measure the relative compensation shares used to weight changes in hours for the various groups of workers.

This section focuses on the development of the weights and hours that enter into the labor composition measure. The main sources of data are the Census of Population, the Current Population Survey (CPS), and the 1973 Social Security Adminis-

[6] For an aggregate measure of labor input to exist, the production function must be separable. See chapter III for a discussion of the production function used in this study.

[7] It is assumed that $g(h_1, h_2, \ldots, h_n)$ is a translog aggregator function which is homogeneous of the first degree. A translog function is a very general form, providing a second-order approximation to any continuous, twice differentiable function. The Tornqvist index in equation 2 is an exact index for the translog function. See W.E. Diewert, "Exact and Superlative Index Numbers," *Journal of Econometrics*, May 1976, pp. 115-145.

[8] This notation is used for simplicity. The growth rates should not be interpreted as instantaneous growth rates but rather as the rate of growth over an interval of time such as a year. Specifically, the measures in this study use the following definition of labor composition change:

$\ln(C_t/C_{t-1}) = \ln(L_t/L_{t-1}) - \ln(H_t/H_{t-1})$.

For small changes, this equation is practically identical to equation 3.

tration exact match file. Work histories and other data from the exact match file form the basis of equations used to estimate the work experience of individuals. The census and CPS data permit wage equations to be estimated with education and estimated work experience as explanatory variables. These procedures are explained below, as is the development of measures of hours of workers. More detailed discussions of these matters appear in appendixes C, D, and E.

The Census and CPS do not collect data on years of work experience. While many researchers have selected age or years since leaving school as a proxy for experience in economic analysis, the approach of this bulletin is to estimate work experience by means of equations. This approach accounts for experience acquired while in school and for interruptions in work histories, which the other measures do not. Accounting for interruptions is particularly important when measuring the work experience of women, who often spend time out of the labor force raising children. Neither age nor years since leaving school captures the effect that raising a family tends to have on work experience.

Experience equations were estimated with data from the 1973 exact match file.[9] The exact match file linked individual records from the Social Security Administration continuous work history files, Internal Revenue Service tax returns, and the March 1973 CPS. The exact match file provided key information regarding the work histories of individuals which was matched to demographic data found in the CPS.

Two experience equations were derived, one for men and one for women. In both cases, the dependent variable was years of work experience reported to the Social Security Administration, including work experience gained while attending school. For men, the explanatory variables included potential experience (age minus years of completed schooling minus six), years of schooling, and interaction terms between potential experience and schooling. The women's equation contained the same variables as well as variables relating to marriage and children. The equations were estimated using linear regression. Appendix C describes the experience equations in detail.

The estimated experience equations generate a measure of work experience for each individual, based on the worker's demographic characteristics. For example, a 35-year-old man who is a college graduate is estimated to have 15 years of work experience. In contrast, a 35-year-old married mother of two children who also graduated from college is estimated to have 9 years of work experience. The considerably lower estimate for such women occurs because most mothers spend some time outside of the labor force, while most male 35-year-olds have been in the labor force continuously.

Once the measures of experience have been estimated, wage equations can be estimated. Wages are used to develop compensation shares in order to weight changes in the hours of each type of worker. As indicated in chapter I, wages can be estimated either as a survey sample average for all persons of a given set of traits or by wage models. Wage rates based on sample averages can be used in constructing a labor composition measure which is broad in scope but less clear in its meaning. Differences in the sample average earnings of workers may reflect not only traits included in the study such as education or experience but also other traits not explicitly included in the study such as part-time work or occupation. Wage models can isolate differentials due to education and work experience from other wage differentials.[10]

This results in a labor composition measure which is more narrow in scope but also more sharply focused on the changes in the composition of the work force by education and work experience. Another advantage of wage equations is that they provide more precise estimates of wages for small groups of workers. Only a few respondents (less than 30) may represent a small group of workers in a survey, and the sample average wage of those few respondents may vary considerably from the average wage of the whole group. With wage equations, the structure of the wage model permits information from all workers to aid in estimating the wage rate for each group of workers. Of course, wage models lead to restrictions on the relationships between the earnings of various groups of workers. These relationships, while examined, tested, and confirmed in hundreds of studies over the last 30 or more years, may not hold exactly.

The wage equations have been estimated with data from the 1950 and 1960 Census of Population and with the Annual Demographic File of the Current Population Survey, 1968 to 1991. The earnings data collected by each survey refer to the preceding year. For example the 1991 CPS collected information on the earnings of workers in 1990. For the years for which the data are not available, linear interpolation of the coefficients of the equations is used to obtain wages.

Earnings data from the decennial censuses and the Current Population Survey are for wage and salary income excluding benefits. For measuring labor composition, total compensation including benefits is preferable. There is no large survey which samples the entire work force, collects demographic characteristics, spans decades, and collects total compensation for workers. Provided that benefits are approximately a constant proportion of total compensation, weights based on earnings will be very similar to weights based on total compensation. Data from the Employment Cost Index (ECI) indicate that the ratio of benefits to total compensation does not

[9] U.S. Department of Health, Education and Welfare, Social Security Administration, "Studies in Interagency Data Linkages: Administrative Record Exact Match File Codebook," Item Pub. No. 79-11750 (Washington, DC, U.S. Government Printing Office, 1979). There are also exact match files for 1980 and 1987, but they are not currently available for use outside of the Census Bureau.

[10] The degree to which wage models can measure just the return to education and experience depends on the other explanatory variables included in the wage model. If other traits are both omitted from the wage model and correlated with education or experience, the education and experience parameters will also include some portion of the return to these other traits. Nevertheless, wage models should generally provide an estimate conceptually closer to the return to education and experience than derived from sample average wage rates.

vary much by occupation.[11] Tests in appendix B with ECI data show little difference in the estimated parameters. Other studies cited in appendix E also suggest that benefits make up a relatively constant proportion of total compensation for most workers.

For each survey year, two equations are estimated, one for men and one for women. Separate equations are used because the pattern of women's earnings differs from that of men, in that the estimated parameters of the wage equations of men and women are significantly different.[12] For each equation, the dependent variable is the natural logarithm of a worker's hourly earnings. The explanatory variables include years of schooling and work experience. Work experience is calculated for each worker from the estimated experience equations. Additional variables such as region of residence also appear in the wage functions to control for earnings differences arising from such factors. Appendix E contains a detailed description of the wage equations including a list of explanatory variables and estimation results for each year.

Tables 12 and 13 show indexes of hourly earnings by education level for men and women, respectively. These measures are constructed with the estimated schooling coefficients of the wage equations. The measures compare the hourly earnings of workers at various education levels to the hourly earnings of high school graduates (that is, those with 12 years of completed schooling). For example, in 1990, a male college graduate (index = 153.4) earned, on average, 53.4 percent more per hour than a male high school graduate (index = 100.0). For women the comparable figure was 53.6 percent. The hourly wage of male high school dropouts (index = 86.5) was 13.5 percent less than the hourly wage of male high school graduates in 1990, while the wage of female high school dropouts (index = 84.1) was 15.9 percent less than the hourly wage of female high school graduates.

Over the past four decades, the hourly wages of those with at least some college education have increased faster than the wages of high school graduates. In 1949, male college graduates made 30.3 percent more per hour than male high school graduates; by 1990 the difference had expanded to 53.4 percent. The relative gains for women are more striking: the differential rose from 10.7 percent in 1949 to 53.6 percent in 1990. Workers with 1 to 3 years of college and workers with postgraduate schooling also improved their positions relative to high school graduates between 1949 and 1990. Those with 1 to 3 years of college have not gained as much as college graduates, whereas those with postgraduate schooling have gained substantially more than college graduates.

Compared to high school graduates, all three of the highest education groups have made gains in the eighties after declines or erratic changes in the seventies. The relative wages of men in these groups first underwent a decline in the first half of the seventies, and the relative earnings of women in those groups moved erratically during the entire decade. In the eighties, the relative earnings of men and women with at least some college generally increased, but workers with more than a college degree made the largest gains.

Table 12. Index of relative hourly earnings of men by education level, 1949-90
(Index, 12 years of completed schooling = 100)

Year	Years of Education						
	1-4	5-8	9-11	12	13-15	16	17+
1949	71.6	84.5	93.7	100.0	109.3	130.3	127.3
1959	71.6	83.6	92.5	100.0	115.6	144.3	145.0
1967	65.0	79.4	87.9	100.0	112.6	133.5	148.5
1968	68.8	78.5	87.8	100.0	113.0	133.4	148.6
1969	63.8	78.4	89.9	100.0	111.2	136.4	149.4
1970	64.4	79.3	88.4	100.0	113.6	132.4	129.6
1971	64.6	78.8	89.7	100.0	113.1	135.3	151.1
1972	64.2	81.2	88.6	100.0	111.0	133.6	143.3
1973	61.5	77.9	90.9	100.0	107.8	126.5	144.0
1974	61.8	80.8	91.4	100.0	106.3	124.3	137.6
1975	57.3	77.2	86.8	100.0	108.0	134.2	149.6
1976	64.9	76.0	85.4	100.0	106.4	133.3	150.8
1977	59.8	73.7	85.9	100.0	107.4	130.5	147.7
1978	57.6	72.8	85.6	100.0	107.8	131.3	147.5
1979	56.7	74.5	85.0	100.0	107.2	128.5	149.9
1980	60.8	74.6	86.0	100.0	106.0	128.9	150.8
1981	57.8	75.5	86.8	100.0	108.0	129.4	149.9
1982	58.1	73.7	83.5	100.0	106.9	129.3	153.8
1983	62.1	74.6	85.6	100.0	108.9	135.6	162.9
1984	59.2	73.1	83.2	100.0	106.6	139.7	159.0
1985	60.2	71.9	86.4	100.0	112.4	139.8	171.8
1986	58.7	72.8	84.0	100.0	110.8	146.1	172.8
1987	63.7	71.4	83.8	100.0	110.4	139.1	171.9
1988	63.3	70.5	84.9	100.0	112.0	145.1	170.4
1989	65.4	72.1	83.2	100.0	115.7	150.7	177.2
1990	65.1	72.8	86.4	100.0	118.2	153.4	178.8

During the postwar period, the hourly wages of workers without a high school diploma have fallen in comparison to the hourly wages of high school graduates. Male and female high school dropouts experienced similar declines. While the wage of male high school dropouts was, on average, 93.7 percent of the wage of male high school graduates in 1949, by 1990 it was down to 86.5 percent. For female high school dropouts, the decline was slightly smaller, from 91.2 percent in 1949 to 84.1 percent in 1990. Those who never attended high school also lost ground relative to high school graduates.

Table 14 presents the percentage change in the wage due to an additional year of experience, which is calculated with the experience coefficients of the wage equations, holding other traits constant. In 1990, a man with 5 years of experience earned 5.9 percent more per hour compared to a man with 4 years of experience. The percentage change in the wage fell as experience rose. At 10 years the gain from an additional year experience was 4.5 percent, while at 15 years it was down to 3.1 percent. At 25 years an additional year of experience led to an increase of just 0.4 percent. The wages for women also grew more slowly as years of experience increased. In 1990

[11] Donald G. Wood, "A New Measure of the Cost of Compensation Components," presented at the Conference on Research in Income and Wealth, May 12-14, 1988, Washington, DC.

[12] Earnings are differentiated by sex in other studies of labor composition, including Jorgenson, Gollop, and Fraumeni, *Productivity and U.S. Economic Growth*; Denison, *Trends in American Economic Growth, 1929-82*; and Chinloy, *Labor Productivity*, ABT Books (Cambridge, MA, 1981). See appendix A for a discussion of earnings differences by sex.

Table 13. Index of relative hourly earnings of women by education level, 1949-90
(Index, 12 years of completed schooling = 100)

Year	Years of Education						
	1-4	5-8	9-11	12	13-15	16	17+
1949	82.0	84.1	91.2	100.0	108.3	110.7	113.3
1959	74.0	83.4	91.1	100.0	110.9	127.1	133.3
1967	74.5	79.7	87.2	100.0	109.5	129.6	140.7
1968	73.1	81.6	88.3	100.0	112.7	125.3	152.5
1969	76.5	84.9	89.1	100.0	113.6	123.0	122.7
1970	71.1	79.5	87.9	100.0	114.0	127.1	109.5
1971	73.0	79.8	89.6	100.0	110.1	128.5	150.7
1972	76.3	81.9	90.5	100.0	111.3	126.2	129.1
1973	76.8	82.4	91.5	100.0	116.0	125.1	140.4
1974	86.0	84.8	89.8	100.0	113.5	123.7	140.9
1975	74.0	80.3	85.0	100.0	112.0	120.6	127.6
1976	75.5	80.5	86.5	100.0	111.7	127.1	143.1
1977	74.1	78.1	86.3	100.0	112.0	127.1	150.7
1978	78.5	80.6	84.3	100.0	110.2	120.3	139.2
1979	73.4	79.4	87.4	100.0	111.3	131.6	145.6
1980	84.2	80.3	87.4	100.0	113.8	132.5	140.2
1981	68.2	81.8	85.2	100.0	115.1	128.5	151.4
1982	62.4	78.4	86.7	100.0	113.1	133.0	149.1
1983	62.4	78.2	85.1	100.0	115.7	133.7	158.4
1984	78.3	79.0	86.4	100.0	117.6	138.0	163.6
1985	67.6	77.1	86.7	100.0	116.1	142.9	165.7
1986	78.3	77.6	83.2	100.0	116.5	145.2	173.7
1987	66.1	73.1	85.6	100.0	118.4	139.9	172.2
1988	66.2	74.4	84.5	100.0	120.5	152.0	175.7
1989	70.0	75.3	83.2	100.0	118.6	155.6	167.1
1990	73.4	73.0	83.7	100.0	118.0	152.8	171.2

an additional year of experience resulted in an increase in the wage of 5.0 percent for women with 5 years of experience, 2.9 percent for women with 10 years of experience and 0.8 percent for women with 15 years of experience. At 25 years of experience the result was a decrease in the wage of 3.2 percent.

The results in Table 14 show that the payoff to additional experience has risen for workers at low and intermediate levels of experience in the postwar period. Most of this increase has occurred since 1970. For example, for men with 1 year of experience, the payoff to an additional year of experience was 5.3 percent in 1949 and 5.4 percent in 1970, while in 1990 it was 7.0 percent. For women with 5 years of experience, the payoff was 2.2 percent in 1949, 2.4 percent in 1970, and 5.0 percent in 1990. This rise in the payoff to experience could be due to increases in the returns to training or to increases in the amount of training per year of experience.

As in table 14, the entries in table 15 were computed from the experience coefficients of the wage equations. Table 15 displays indexes of relative earnings by experience level for men and women for six levels of experience (0, 5, 10, 15, 25, and 30 years). In each half of the table, all explanatory variables are held constant except experience. For instance, in 1990 a man with 10 years of experience earned 75.0 percent more per hour than a man with no experience who was identical in other respects. With 25 years of experience, his wage would have been 144.8 percent higher than the wage of the man who had no experience.

Table 14. Percent change in wages due to an additional year of work experience, men and women, selected years, 1949-90[1]

Year	Years of work experience					
	1	5	10	15	25	30
Men						
1949	5.3	4.4	3.3	2.2	-0.1	-1.1
1959	5.9	4.9	3.7	2.5	.1	-1.1
1970	5.4	4.5	3.3	2.1	-.2	-1.3
1975	7.9	6.4	4.7	2.9	-.5	-2.1
1980	7.7	6.3	4.7	3.0	-.2	-1.7
1985	8.4	6.9	5.2	3.5	.2	-1.5
1990	7.0	5.9	4.5	3.1	.4	-1.0
Women						
1949	2.8	2.2	1.4	0.6	-0.9	-1.6
1959	2.7	2.3	1.7	1.1	-.1	-.7
1970	3.2	2.4	1.5	.5	-1.3	-2.2
1975	5.0	3.8	2.2	.6	-2.4	-3.9
1980	4.9	3.7	2.2	.8	-2.3	-3.8
1985	6.4	4.8	2.9	.9	-2.8	-4.6
1990	6.7	5.0	2.9	.8	-3.2	-5.1

[1] Measured as the percentage difference in hourly wages between a worker with a given number of years of work experience and 1 year less of experience, for example, the percentage change in hourly wages between 4 and 5 years of work experience.

Charts 3 and 4 indicate that the earnings-experience profiles have become steeper during the past 40 years. In 1949 a man with 15 years of experience earned 73.6 percent more per hour than he would have with no experience, while in 1990 the differential was up to 109.4 percent. A woman with 15 years of experience in 1949 earned a wage 29.1 percent higher than a woman with no experience, while in 1990 her wage was 72.9 percent higher.

Table 15. Index of relative hourly earnings by years of work experience, selected years, 1949-90
(Index, 0 years of work experience = 100)

Year	Years of work experience					
	0	5	10	15	25	30
Men						
1949	100.0	126.9	152.6	173.6	190.6	184.0
1959	100.0	130.4	160.1	185.2	207.4	200.7
1970	100.0	127.3	153.1	174.0	189.2	181.1
1975	100.0	141.3	183.6	219.4	243.2	225.8
1980	100.0	140.2	181.6	217.4	246.1	232.6
1985	100.0	144.6	192.5	236.2	278.0	266.8
1990	100.0	136.8	175.0	209.4	244.8	239.3
Women						
1949	100.0	113.1	123.1	129.1	126.7	118.7
1959	100.0	113.2	124.3	132.7	138.3	135.1
1970	100.0	114.9	126.0	131.9	125.8	114.6
1975	100.0	124.0	142.5	151.6	136.3	115.2
1980	100.0	123.5	141.5	150.5	136.3	115.9
1985	100.0	131.5	157.3	171.2	152.6	125.0
1990	100.0	132.8	159.4	172.9	150.2	120.3

The most striking difference between the men's and women's profiles is that the women's profile is much flatter throughout the period (chart 5). In 1949 a woman with 10 years of experience earned 23.1 percent more per hour than if

Chart 3. Indexes of hourly earnings of men by years of work experience, selected years, 1948-90

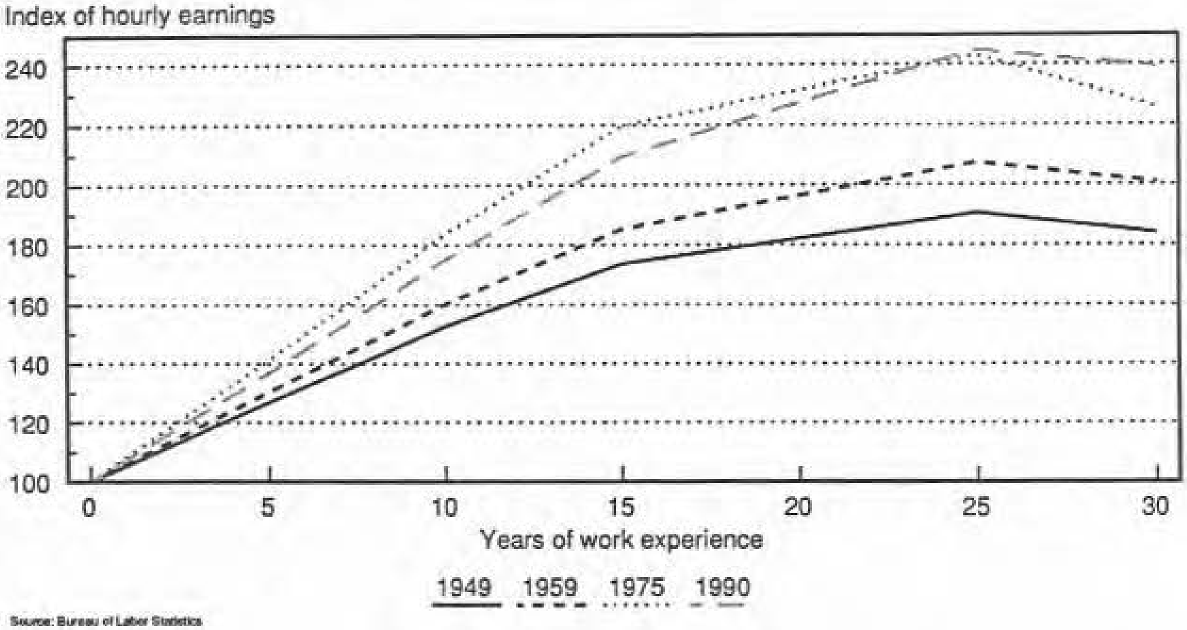


Chart 4. Indexes of hourly earnings of women by years of work experience, selected years, 1948-90

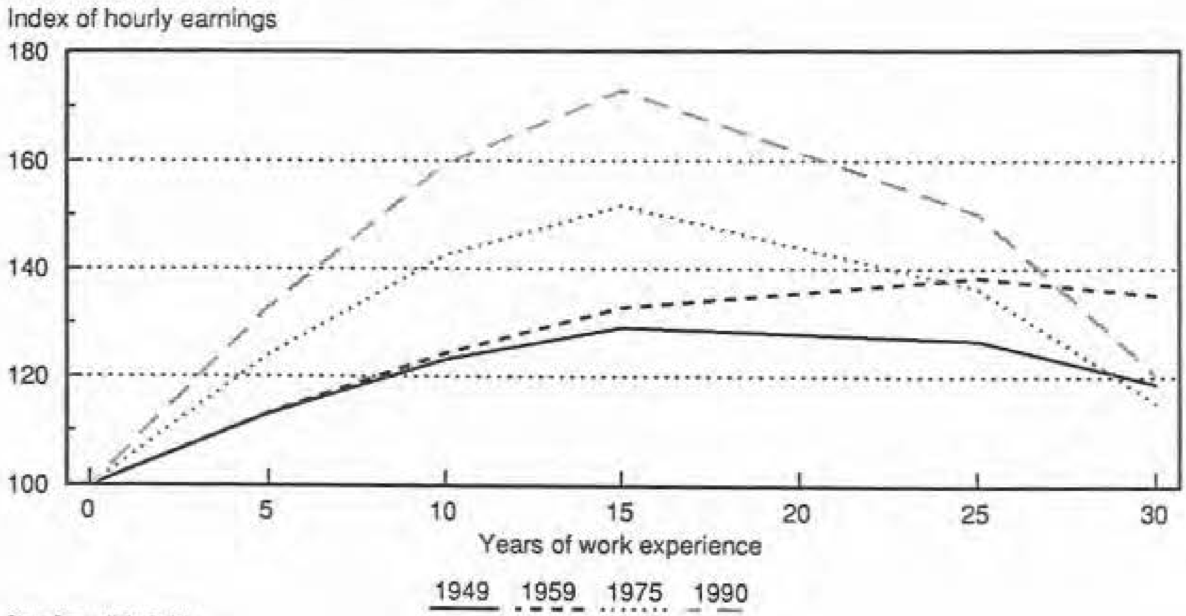
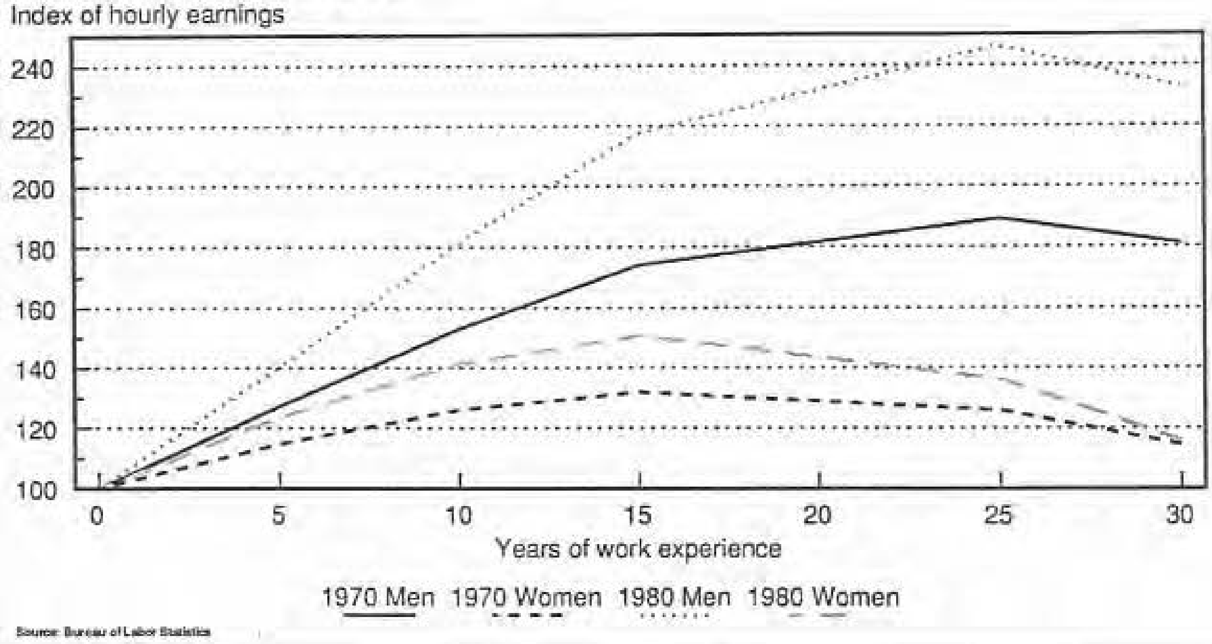


Chart 5. Indexes of hourly earnings of men and women by years of work experience, selected years, 1948-90



she had had no experience, while for men the differential was 52.6 percent. Forty-one years later, a woman with 10 years of experience made 59.4 percent more per hour than if she had had no experience, while for men the figure was 75.0 percent.

In addition to information on earnings, data on annual hours of work are needed to measure labor composition growth. Hours matrices have been created with data from the 1950 and 1960 Census of Population and the Annual Demographic File of the Current Population Survey, 1968 to 1991. (As with the earnings data, the annual hours data collected by a survey refer to the previous year.) For the remaining years, an iterative multiproportional interpolation method was used to complete the set of matrices for the period 1948 to 1990.[13] For this set of matrices, hours of work are cross-classified by age, education, and sex; the hours of female workers are also classified by marital status and number of children ever born. The characteristics and dimensions of the matrices are presented in table 16. In total, there are 4,032 cells for women and 504 for men. The experience equations are used to condense these matrices into a set of annual hours matrices which are cross-classified by estimated work experience, education, and sex. These smaller matrices each have 504 cells.[14] Each cell of each matrix contains the total number of hours worked in that year by a particular group of workers; all members within a cell are the same sex and have the same amount of education and work experience. Labor composition growth is

[13] See M. Bacharach, "Estimating Nonnegative Matrices from Marginal Data," *International Economic Review*, September 1965, pp. 294-310.

[14] There are 72 levels of estimated experience (0 to 71 years) and 7 levels of education for each sex.

Table 16. Characteristics and dimensions of annual hours matrices

Characteristic	Dimension		Types
	Men	Women	
Education	7	7	0-4 years of schooling 5-8 9-11 12 13-15 16 17 or more
Age	72	72	14-85 years old
Marital status	-	2	Ever married Never married
Number of children ..	-	4	None 1 2-3 4 or more
Total cells	504	4,032	

computed with this set of matrices. See appendix D for details concerning the construction of the hours matrices.

Labor input and labor composition growth

Once the hours of labor for each type of worker are collected and the corresponding wage rate estimated, changes in composition-adjusted labor input are measured. Changes in labor input are measured using the Tornqvist index formula of equation 2 and then linked, forming the Tornqvist index of labor input in the private business sector as shown in table 17.

Between 1948 and 1990, labor input grew 71.3 percent or 1.3 percent annually in private business. The rate of growth of labor input increased greatly after 1973. Making comparisons from one peak of the business cycle to another reveals the trend in labor input growth with a minimum of cyclical fluctuations in the demand for labor. From 1948 to 1973, labor input grew at an average annual rate of 0.9 percent. Since 1973, labor input has increased 1.9 percent annually. During this later period labor input has grown at a fairly consistent rate. During the period 1973-79, labor input grew 1.8 percent annually. Since 1979, labor input has risen at a slightly higher rate, 2.0 percent.

Table 18 contains the Tornqvist index of labor input for private nonfarm business. Labor input in this sector, which excludes farm production, grew 31 percentage points more than in the private business sector primarily due to a shift in employment from farm to nonfarm industries during the 1950's and 1960's. Labor input grew 103.1 percent after 1948 or 1.7 percent annually. Again, labor input growth proceeded at a faster rate after 1973 than during the 1948-73 period. In both sectors, labor input grew faster than total hours. Since 1948, labor input in the private business and private nonfarm business sectors increased at average annual rates of 1.3 and 1.7 percent, respectively, while hours increased only 1.0 and 1.4 percent.

The difference between the growth in labor input and total hours equals the change in the composition of the work force. Between 1948 and 1990, labor composition growth resulting from a rising level of worker skills added 13.3 percent to labor input in private business and 12.0 percent in private nonfarm business. Labor composition in private business grew 0.30 percent annually from 1948 to 1990 (table 19). Labor composition in private nonfarm business increased almost as rapidly, at 0.27 percent annually.

Between 1948 and 1990, more than 75 percent of the growth in private business labor input came from the increase in the hours of labor, and the remaining 25 percent came from labor composition. Prior to 1960, however, chart 6 indicates that all of the growth of labor input was due to labor composition growth. In the next two periods shown, labor composition growth never accounted for more than 25 percent of the growth in labor input. In addition, for the period 1960-79 labor composition growth accounted for just 8 percent of the increase in labor input. Although the growth rate of labor composition may seem small, especially when compared to the

Chart 6. Contribution of hours and labor composition to labor input growth in private business, 1948-90

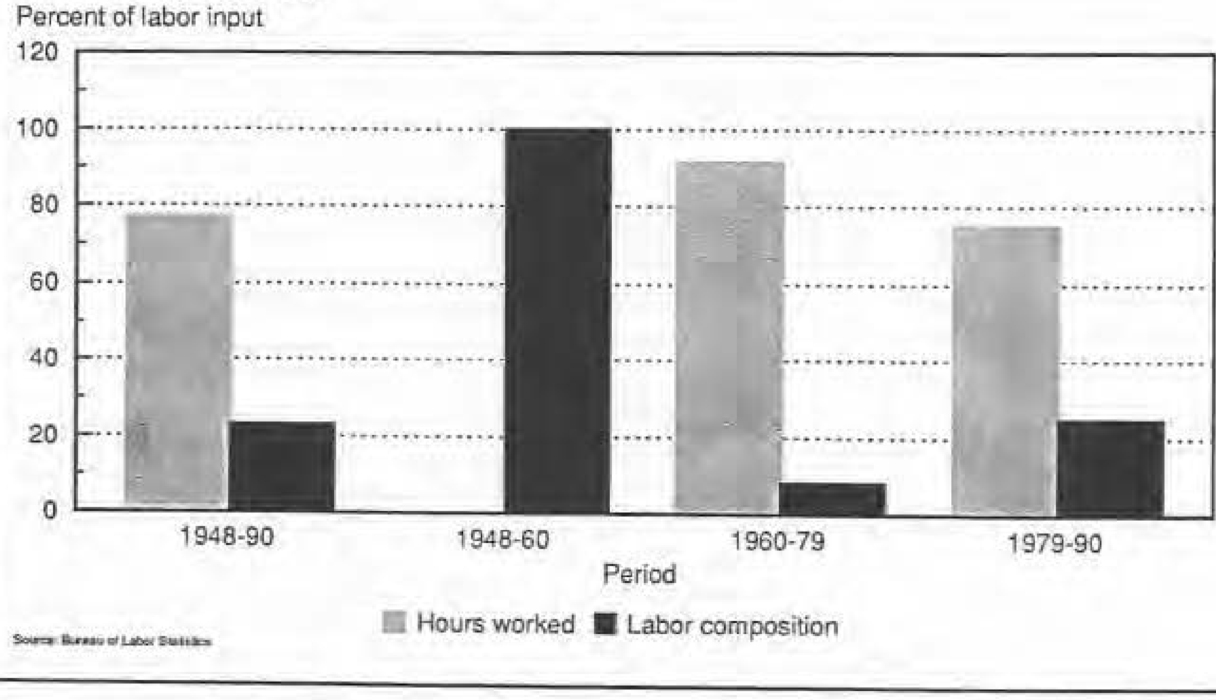


Table 17. Index of labor input and rate of change in the U.S. private business sector, 1948-90
(Index 1982 = 100)

Year	Tornqvist index of labor input	Percent change
1948	73.02	-
1949	71.16	-2.55
1950	71.89	1.03
1951	74.12	3.10
1952	74.23	.15
1953	75.43	1.62
1954	73.45	-2.62
1955	76.32	3.91
1956	77.58	1.65
1957	76.59	-1.28
1958	73.27	-4.33
1959	76.22	4.03
1960	76.61	.51
1961	75.78	-1.08
1962	77.66	2.48
1963	78.23	.73
1964	79.43	1.53
1965	81.89	3.10
1966	83.72	2.23
1967	83.56	-.19
1968	84.54	1.17
1969	86.91	2.80
1970	85.53	-1.59
1971	84.89	-.75
1972	87.77	3.39
1973	90.70	3.34
1974	91.31	.67
1975	87.42	-4.26
1976	89.78	2.70
1977	93.31	3.93
1978	97.95	4.76
1979	100.76	2.87
1980	100.10	-.66
1981	101.56	1.46
1982	100.00	-1.54
1983	102.24	2.24
1984	108.28	5.91
1985	110.77	2.30
1986	112.03	1.14
1987	115.77	3.34
1988	120.69	4.25
1989	124.35	3.03
1990	124.98	.51
Average annual growth rates		
1948-90		1.29
1948-73		.87
1973-90		1.90
1973-79		1.77
1979-90		1.98

growth of total hours, increases in the skill level of the work force contributed about 9 percentage points to the growth in private business sector output after 1948.[15]

[15] In this study, only the direct contribution of increased skills to labor input is considered. Better training and higher educational attainment may raise output through indirect means as well. For example, Finis Welch, "Education in Production," *Journal of Political Economy*, 78, no. 1, January/February 1970, pp. 35-59, finds that more highly educated farmers adopt new technology more quickly. A. Bartel and F. Lichtenberg, "The Comparative Advantage of Educated Workers in Implementing New Technology," *Review of Economics and Statistics*, 66, no. 1, February 1987, pp. 1-17, argue that more highly educated workers have a comparative cost advantage in implementing new technology. If wages do not fully reflect these skills, labor composition growth will be understated.

Unlike total labor input, labor composition in the private business sector increased at only a slightly faster pace after 1973 (table 19). A somewhat similar pattern is found for the nonfarm private business sector. On closer inspection, however, one sees that labor composition changed at distinctly different rates over the course of the last 42 years. Labor composition growth was most rapid in the 1950's and 1980's but very slow in the 1960's and 1970's. Labor composition grew 0.40 percent annually from 1948 to 1960 in private business. Similarly, the average annual growth rate has been 0.49 percent since 1979. During the 1960-79 period, labor composition was virtually unchanged, increasing at an annual rate of only 0.12 percent.

The pattern of labor composition growth results largely from the changes in the underlying experience distribution of the work force. As indicated earlier, educational attainment has grown at a more or less steady rate over the past 42 years and therefore is not a major factor in the fluctuation of labor composition growth. However, the average level of work experience began to decline around 1959. The experience level of all workers continued to fall until 1979, and it has increased slightly in the 1980's. The average level of experience declined primarily because the baby boom generation entered the work force. In addition, the relatively low youth unemployment rates of the 1960's and early 1970's may indicate a scarcity of labor which may have forced employers to hire less experienced workers. Furthermore, the large increase in labor force participation of women during this period may have been primarily among young women and re-entrants with low levels of previous work experience, thus causing further declines in the average level of work experience.[16]

Labor composition growth also tends to move counter to the business cycle. Usually, labor composition growth is faster during economic downturns and slower during expansions when the demand for labor approaches its peak. Labor composition in the private nonfarm business sector grew about 50 percent faster in the 11 years when total hours decreased than in the 31 years when hours increased.

Average percentage change in private nonfarm business labor composition by change in hours, 1949-90
(Percentage change in labor composition)

All years	0.27
Years when total hours rise	.24
Years when total hours fall	.36

If firms lay off their least productive workers first during an economic decline, the average skill level of the remaining work force will increase. Since firms often follow a "last hired, first fired" scheme, this hiring practice contributes to an increase in the average level of experience of the work force in periods of declining labor demand. Furthermore, highly educated workers are less likely to be laid off during a reces-

[16] See June O'Neill, "The Trend in the Male-Female Wage Gap in the United States," *Journal of Labor Economics*, pt. 2 supp. January 1985, pp. S91-116.

Table 18. Index of labor input and rate of change in the U.S. private nonfarm business sector, 1948-90
(Index 1982 = 100)

Year	Tornqvist index of labor input	Percent change
1948	62.55	-
1949	60.14	-3.85
1950	62.29	3.57
1951	65.18	4.64
1952	66.13	1.46
1953	68.34	3.34
1954	66.35	-2.91
1955	69.25	4.37
1956	71.07	2.63
1957	70.97	-.14
1958	68.00	-4.18
1959	71.05	4.49
1960	71.44	.55
1961	71.25	-.27
1962	73.45	3.09
1963	74.35	1.23
1964	76.08	2.33
1965	78.83	3.61
1966	81.31	3.15
1967	81.27	-.05
1968	82.30	1.27
1969	85.10	3.40
1970	83.97	-1.33
1971	83.54	-.51
1972	86.35	3.36
1973	89.59	3.75
1974	90.20	.68
1975	86.36	-4.26
1976	88.95	3.00
1977	92.73	4.25
1978	97.50	5.14
1979	100.51	3.09
1980	99.96	-.55
1981	101.52	1.56
1982	100.00	-1.50
1983	102.38	2.38
1984	108.67	6.14
1985	111.56	2.66
1986	113.12	1.40
1987	117.09	3.51
1988	122.23	4.39
1989	126.21	3.26
1990	126.97	.60
Average annual growth rates		
1948-90		1.70
1948-73		1.45
1973-90		2.07
1973-79		1.94
1979-90		2.15

sion. Other factors being equal, labor composition can be expected to rise more rapidly during an economic downturn. During an expansion, the reverse pattern is observed. Firms generally recall their best workers first and then recall the more marginal workers as the expansion continues. At the peak, firms consider hiring new workers, some of whom may not meet previous hiring standards. As the expansion proceeds, labor composition increases less and less.[17]

Table 20 compares the growth rates of three different measures of labor input in private nonfarm business: Employment, hours, and the Tornqvist index of labor input. In each period shown, employment grew more rapidly than hours, because average weekly hours were falling. Also, the Tornqvist index of labor input rose faster than hours in each period, be-

[17] In the year preceding a fall in total hours, labor composition in the private nonfarm business sector grew at about 75 percent of the average growth rate.

Table 19. Index of labor composition and rate of change in the U.S. private business and private nonfarm business sectors, 1948-90
(Index 1982 = 100)

Year	Private business		Private nonfarm business	
	Labor composition	Percent change	Labor composition	Percent change
1948	91.27	-	92.25	-
1949	91.47	0.21	92.38	0.14
1950	92.17	.77	92.83	.49
1951	92.41	.26	92.85	.02
1952	92.67	.27	93.53	.74
1953	93.36	.75	94.27	.78
1954	94.04	.73	94.92	.69
1955	94.22	.19	95.13	.22
1956	94.37	.17	95.27	.15
1957	94.78	.43	95.78	.53
1958	95.15	.39	95.92	.14
1959	95.27	.12	96.01	.10
1960	95.76	.51	96.02	.01
1961	96.29	.56	96.94	.96
1962	97.20	.94	97.93	1.02
1963	97.42	.22	98.09	.16
1964	97.46	.05	98.04	-.05
1965	97.37	-.10	97.93	-.11
1966	97.35	-.02	97.72	-.21
1967	97.50	.16	97.68	-.04
1968	97.28	-.23	97.39	-.30
1969	97.65	.38	97.81	.43
1970	98.08	.44	98.21	.41
1971	97.80	-.29	97.94	-.28
1972	97.85	.05	97.90	-.04
1973	97.63	-.22	97.69	-.21
1974	98.28	.67	98.36	.68
1975	98.33	.05	98.36	.00
1976	98.12	-.22	98.08	-.29
1977	98.12	.00	98.12	.05
1978	98.24	.13	98.18	.06
1979	97.92	-.33	97.87	-.32
1980	98.24	.32	98.19	.33
1981	98.99	.76	98.95	.77
1982	100.00	1.02	100.00	1.07
1983	100.43	.43	100.37	.37
1984	100.54	.11	100.43	.06
1985	100.79	.25	100.68	.25
1986	101.30	.50	101.27	.58
1987	101.55	.25	101.47	.19
1988	102.37	.81	102.29	.81
1989	102.86	.47	102.77	.46
1990	103.37	.51	103.31	.52
Average annual growth rates				
1948-90		.30		.27
1948-73		.27		.23
1973-90		.34		.33
1973-79		.05		.03
1979-90		.49		.49
1948-60		.40		.33
1960-79		.12		.10
1979-90		.49		.49

cause of increases in labor composition. In most of the periods, the growth rate of the Tornqvist index was below the growth rate of employment, implying that the drop in average weekly hours was greater than the rise in labor composition. The exception is the interval between 1979 and 1990, when the Tornqvist index of labor input increased faster than employment.

The importance of considering average weekly hours and labor composition when measuring labor input is illustrated by the last two rows of table 20. Using employment to gauge the change in labor input, one would conclude that labor input grew much more quickly in 1973-79 than in 1979-90; employment rose at average annual rates of 2.64 percent and 1.82 percent in those time spans, respectively. However, average weekly hours fell more rapidly in the earlier span, and labor composition advanced with greater speed in the later span. The net effect was that, based on the Tornqvist index, labor input actually grew more rapidly between 1979 and 1990 than it did between 1973 and 1979.

Table 20. Employment, hours, and the Tornqvist index of labor input in private nonfarm business, 1948-90
(percentage change)

Years	Average annual growth rates		
	Employment	Hours	Tornqvist index of labor input
1948-90	1.80	1.43	1.70
1948-73	1.59	1.22	1.45
1973-90	2.11	1.74	2.07
1973-79	2.64	1.91	1.94
1979-90	1.82	1.65	2.15

Note: The differences between the growth rates of the Tornqvist index of labor input and hours are not all exactly equal to the growth rates of labor composition in table 17 due to rounding.

Chapter III. U.S. Productivity Growth, 1948-90

The measurement of labor composition is only the first step toward assessing its contribution to productivity growth in the U.S. economy. This chapter presents a growth accounting framework for analyzing the contribution of labor composition to output per hour (labor productivity). The chapter begins by reviewing trends in labor productivity in the postwar era. A framework for analyzing labor productivity growth is then presented. Next, the effect of labor composition on productivity is assessed. The analysis then expands to include the role of capital services, and new BLS measures of multifactor productivity are introduced.

Labor productivity trends, 1948-90

The business sector is the broadest segment of the economy for which the Bureau of Labor Statistics measures labor productivity. Business sector output is defined as gross domestic product (GDP) in constant 1982 dollars less the output of general government, nonprofit institutions, paid employees of private households, the rental value of owner-occupied dwellings, and the statistical discrepancy in computing the National Income and Product Accounts. Some of these items are excluded because output measures for them are relatively weak, being defined, in some cases, in terms of labor compensation. Other items are excluded because there are no corresponding measures of factors of production.[1] The business sector has averaged about 78 percent of GDP in recent years. Nonfarm business also excludes the output of the farm sector and has averaged about 77 percent of GDP in recent years.

Chart 7 shows rising levels of output, hours at work, and labor productivity, defined as output per hour at work, for the business sector. Since 1948, business sector output has grown 281.2 percent, while hours increased 52.9 percent. As a result, labor productivity grew 148.9 percent or at an annual rate of 2.2 percent.

Labor productivity in business has consistently risen in all postwar expansions. However, during recessions labor productivity has grown only slightly or has declined. These cyclical changes in labor productivity arise from larger variations in the growth rate of output compared to labor. Changes in the utilization of the capital stock contribute to the cyclical variations in output growth rates. Cyclical variations in labor may be dampened by lags in hiring and firing workers or by the hoarding of workers with specialized skills.[2]

Comparing productivity levels between peaks of the business cycle can highlight long-term trends in productivity growth. Table 21 displays the average annual growth rates of labor productivity in the business sector between business cycle peaks. With just one exception, labor productivity growth rates exceeded 2.0 percent in all periods prior to 1973 for both the business and nonfarm business sectors. After 1973 the rates of growth were very much slower. In the three completed business cycles since 1973, labor productivity has not achieved the growth rates recorded for the five previous business cycles.

The slowdown in labor productivity has been the subject of intense investigation.[3] Though some researchers mark the beginning of the productivity slowdown as early as 1967, the widely accepted starting point of the slowdown is 1973. In private business (which excludes government enterprises), labor productivity has been growing 0.9 percent per year since 1973 (table 22). This rate is less than a third of the 3.1 percent growth rate witnessed between 1948 and 1973. Accompanying the slower growth in productivity is a slower rate of growth in output and a faster rate of growth in hours at

[1] See Jerome Mark, "Problems Encountered in Measuring Single- and Multifactor Productivity," *Monthly Labor Review*, December 1986, pp. 3-11, for a discussion of measurement issues.

[2] Walter Oi, "Labor as a Quasi-Fixed Factor," *Journal of Political Economy*, December 1962, pp. 538-55, demonstrates that fixed costs associated with employing labor means that temporary changes in output lead to smaller changes in labor.

Jon Fay and James Medoff, "Labor and Output Over the Business Cycle: Some Direct Evidence," *American Economic Review*, June 1985, pp. 638-655, estimate that about 5 percent of labor is hoarded during the trough quarter of the business cycle.

See Susan Powers, "The Role of Capital Discards in Multifactor Productivity Measurement," *Monthly Labor Review*, June 1988, and "Cyclical Variation in Capital Stock Measures: Implications for Multifactor Productivity," Ph.D thesis (State University of New York at Binghamton, 1985) for a discussion of cyclical influences on measures of capital.

[3] For a discussion of potential causes of the productivity slowdown, see Edwin Dean and Kent Kunze, "Recent Changes in the Growth of U.S. Multifactor Productivity," *Monthly Labor Review*, May 1988, pp. 14-22; Edward Denison, *Trends in American Economic Growth, 1929-1982*, The Brookings Institution (Washington, DC, 1985); Martin N. Baily and Robert Gordon, "Measurement Issues, the Productivity Slowdown, and the Explosion of Computer Power," *Brookings Papers on Economic Activity*, 2, 1988, pp. 347-420; J. Randolph Norsworthy, Michael Harper, and Kent Kunze, "The Slowdown in Productivity Growth: Analysis of Some Contributing Factors," *Brookings Papers on Economic Activity*, 2, 1979, pp. 387-421; and the U.S. Department of Labor, Bureau of Labor Statistics, *Trends in Multifactor Productivity, 1948-81*, Bulletin 2178, U.S. Government Printing Office (Washington, DC, 1983).

Table 21. Average annual growth rates of output per hour in business and nonfarm business between business cycle peaks, 1948-90[1]
(Percent per year)

Period	Business[2]	Nonfarm business[2]
1948 IV - 1953 III	4.2	3.1
1953 III - 1957 III	2.0	1.6
1957 III - 1960 II	2.9	2.3
1960 II - 1969 IV	3.0	2.5
1969 IV - 1973 IV	2.3	2.1
1973 IV - 1980 I	.7	.6
1980 I - 1981 III	1.3	.8
1981 III - 1990 III	1.2	1.1

[1] Cyclical peaks are those designated by the National Bureau of Economic Research.

[2] Includes government enterprises.

work. So, in recent years, smaller increases in output have been achieved with larger increases in hours.

Although the 2.2-percent slowdown in the growth rate of private business labor productivity may appear small, the cumulative effect is quite large. If no slowdown had occurred, labor productivity would have been 44 percent higher in 1990. Seen in a more dramatic light, all other things unchanged, real private business sector output in 1990 would have been more than $1.4 trillion higher (1982 dollars) without a productivity slowdown.

Table 22. Average annual rates of growth of output per hour in private business and private nonfarm business for selected periods, 1948-90[1]
(Percent per year)

Period	Output per hour	Output	Hours
Private business			
1948-90	2.2	3.2	1.0
1948-73	3.1	3.7	.6
1973-90	.9	2.4	1.6
1973-79	.8	2.5	1.6
1979-90	.9	2.4	1.6
Private nonfarm business			
1948-90	1.8	3.2	1.4
1948-73	2.5	3.8	1.2
1973-90	.7	2.4	1.7
1973-79	.6	2.5	1.9
1979-90	.7	2.4	1.6

[1] Excludes government enterprises.

Chart 8 shows the course of labor productivity in nonfarm business since 1948. Labor productivity grew 114.1 percent over the period or at an annual growth rate of 1.8 percent. Labor productivity in nonfarm business grew more slowly than in business because farm sector productivity grew very rapidly, especially during the 1950's and 1960's.

The productivity slowdown was almost as large in private nonfarm business. Prior to 1973, labor productivity grew at an annual rate of 2.5 percent, but only 0.7 percent from 1973 to 1990 (table 22). Again, the 1.8-percent reduction in the growth rate of labor productivity represents a large decline in potential output. If no slowdown had occurred, labor productivity would have been 36 percent higher, and real private nonfarm business output would have been more than $1.1 trillion higher (1982 dollars) in 1990.

Measuring labor productivity changes

A growth accounting framework relates changes in labor productivity to changes in outputs and the inputs used in production. This framework explicitly shows how changes in the composition of the work force can account for changes in labor productivity. The derivation of the growth accounting framework may begin with a production function which relates the quantity of output or value added, Q, which can be produced with capital services (for simplicity capital is not disaggregated), k, many types of labor inputs, $h_1, \ldots, h_n$, and the technology available at time t, A_t.[4]

$$1)\ Q = A_t * f(k, h_1, \ldots, h_i, \ldots, h_n)$$

Differentiating the logarithm of equation 1 with respect to time yields an equation for the growth rate of output. The boldface letters represent the growth rate of the variables, for example, **q** is the growth rate of output, Q.[5]

$$2)\ \mathbf{q} = \mathbf{a} + s_k * \mathbf{k} + s_{h1} * \mathbf{h}_1 + \ldots + s_{hi} * \mathbf{h}_i + \ldots + s_{hn} * \mathbf{h}_n$$

The growth rate of output, **q**, depends on the growth rate of capital services, **k**, the growth rate of each type of labor services, **h**, and the growth rate of multifactor productivity, **a**. The terms, s_k and s_{hi}, are the output elasticities of capital and each type of labor, that is, the percentage changes in output due to a 1-percent increase in a given type of capital or labor, respectively.

Multifactor productivity measures the increase in output which cannot be attributed to increases in labor *or* capital. In other words, multifactor productivity measures the change in output arising from changes in technology, the utilization of capital, organization of production, managerial skills, and investments in research and development. By rearranging terms, multifactor productivity can be seen to be the growth rate of output less the weighted growth rates of capital and labor.

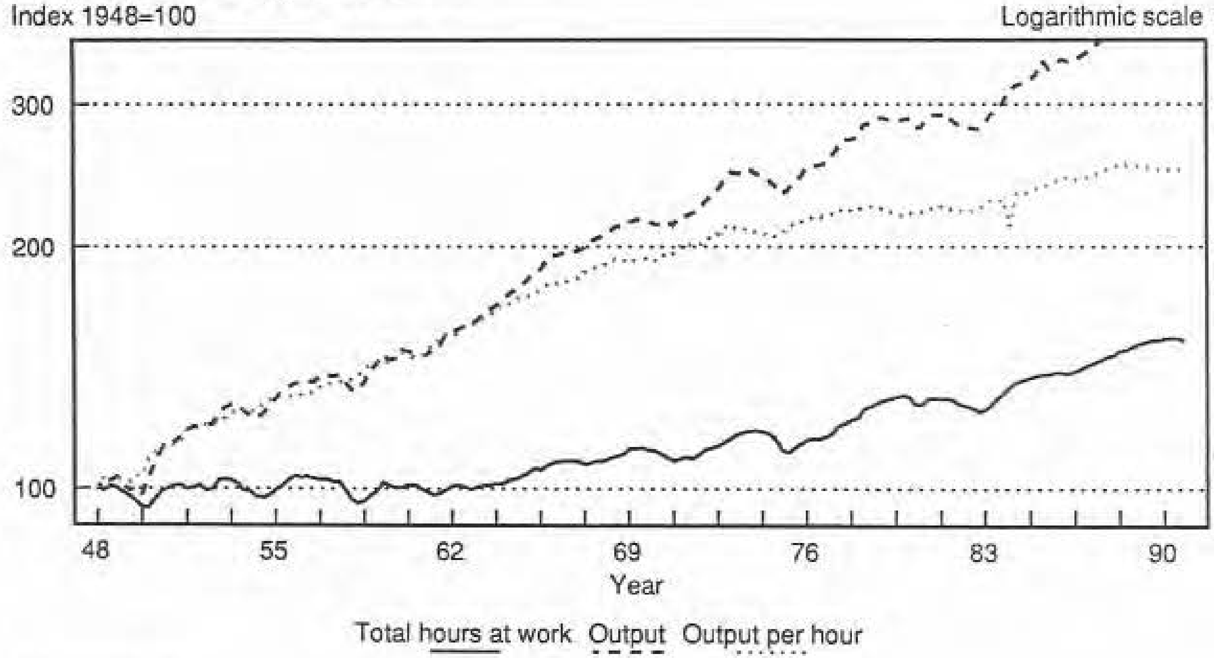
$$3)\ \mathbf{a} = \mathbf{q} - s_k * \mathbf{k} - s_{h1} * \mathbf{h}_1 - \ldots - s_{hi} * \mathbf{h}_i - \ldots - s_{hn} * \mathbf{h}_n$$

By assuming that labor is an input separable from capital, it is possible to concentrate exclusively on an aggregate of labor

[4] Multifactor productivity is derived in appendix A using a more general description of the production function. However, meaningful measures of labor input and composition require the Hicks' neutral technical change used in equation 1.

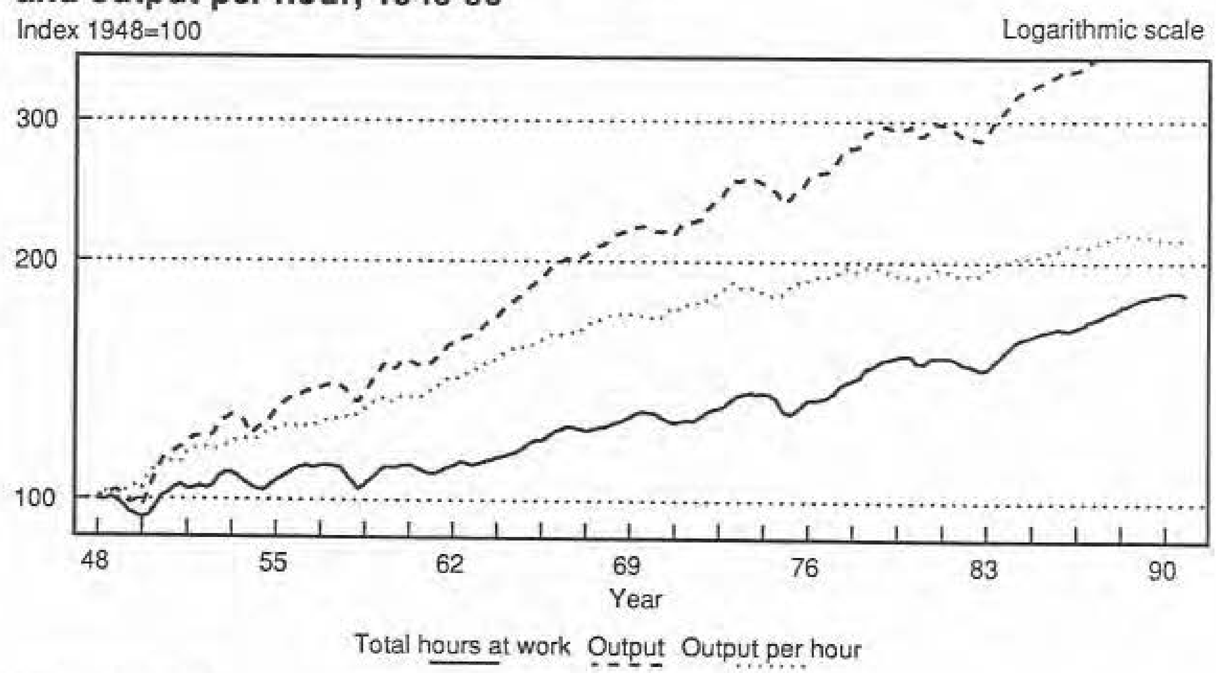
[5] This notation is used for simplicity. The growth rates should not be interpreted as instantaneous growth rates but rather as the rate of growth over an interval of time such as a year. Again, appendix A provides a more rigorous derivation of multifactor productivity and labor composition.

Chart 7. Business output, hours, and output per hour, 1948-90



Source: Bureau of Labor Statistics

Chart 8. Nonfarm business output, hours, and output per hour, 1948-90



Source: Bureau of Labor Statistics

input.[6] Equation 3 can be rewritten in terms of aggregate labor input, **l**, and labor's share of total costs s_l.

4) $\mathbf{a} = \mathbf{q} - s_k * \mathbf{k} - s_l * \mathbf{l}$

Aggregate labor input, **l**, is then a weighted average of each type of hours where the weights, s_{li} $(= s_{hi}/s_l)$, are each type of worker's share of labor compensation. Recall that the growth rate of labor inputs, **l**, is the sum of the growth rate of total hours, **H**, and labor composition, **c**.

5) $\mathbf{l} = s_{l1} * \mathbf{h}_1 + \ldots + s_{hi} * \mathbf{h}_i + \ldots + s_{ln} * \mathbf{h}_n = \mathbf{H} + \mathbf{c}$

Substituting equation 5 into equation 4 yields another expression for multifactor productivity growth:

6) $\mathbf{a} = \mathbf{q} - s_k * \mathbf{k} - s_l * (\mathbf{H} + \mathbf{c})$

The growth rate of labor productivity is defined as the growth rate of output less the growth rate of total hours at work. Rearranging the terms of equation 6 and subtracting the growth rate of hours from each side yields the result:

7) $\mathbf{q} - \mathbf{H} = \mathbf{a} + s_k * \mathbf{k} + (s_l - 1) * \mathbf{H} + s_l * \mathbf{c}$

If industry is perfectly competitive in both the product and factor markets and if the production function results in constant returns to scale, the output elasticities of capital and labor equal their shares of total cost, and the sum of these two shares is one. So, $s_k = (1 - s_l)$ and equation 7 can be rewritten:

8) $\mathbf{q} - \mathbf{H} = \mathbf{a} + s_k * (\mathbf{k} - \mathbf{H}) + s_l * \mathbf{c}$

Equation 8 shows that labor productivity and multifactor productivity are related measures. The growth rate of labor productivity is the sum of the growth rates of multifactor productivity, the contribution of capital intensity (the growth rate of capital less hours at work weighted by capital's cost share), and the contribution of labor composition (the growth rate of labor composition weighted by labor's cost share).

Labor productivity change thus results from two sources of growth not included in multifactor productivity, changes in capital intensity and changes in labor composition. Provided the quantity of capital per hour and the skill level of the work force remains unchanged, labor productivity growth will equal multifactor productivity growth. As will be seen, both capital intensity and labor composition have been increasing over the last 42 years, and so labor productivity has been growing at a faster pace than multifactor productivity.

The measure of multifactor productivity in equation 8 differs from earlier BLS measures (see *Trends in Multifactor Productivity, 1948-81*). Previous multifactor productivity measures, **b**, assumed a homogeneous labor input rather than a Tornqvist index of labor input, and its relationship to labor productivity is slightly different than that found in equation 8.

9) $\mathbf{q} - \mathbf{H} = \mathbf{b} + s_k (\mathbf{k} - \mathbf{H})$

Equation 9 implies that measures of multifactor productivity change based on homogeneous hours, **b**, include the effect of changes in labor composition. The measures of multifactor productivity based on weighted labor input in equation 8, **a**, treat changes in labor composition as a separate effect, and so the measures of multifactor productivity presented in this bulletin exclude the effect of labor composition. Formally, the growth rate of the weighted-labor-based measure of multifactor productivity, **a**, is the difference between the growth rates of the homogeneous-labor-based measure of multifactor productivity, **b**, and the contribution of labor composition, $s_l * \mathbf{c}$.

10) $\mathbf{a} = \mathbf{b} - s_l * \mathbf{c}$

Sources of labor productivity growth in private business

Equation 8 provides an accounting framework for identifying sources of labor productivity growth. Since there are no measures of capital input for government enterprises, these enterprises are excluded from both the labor and multifactor productivity measures. As such, private business is the broadest sector available for consistent measurement of both labor and multifactor productivity. For the entire period 1948-90, labor productivity increased 2.2 percent annually in private business. Table 23 shows the annual contributions of changes in labor composition, capital intensity, and multifactor productivity to changes in labor productivity in private business.[7]

Changes in the education, work experience, and sex composition of the work force have contributed modestly to labor productivity. From 1948 to 1990, shifts in the composition of the work force contributed 0.2 percentage point to the 2.2 percent annual average growth rate of labor productivity in the private business sector. On average, then, labor composition accounts for 9 percent of the increase in labor productivity.

Table 23. Sources of labor productivity growth in private business, 1948-90 and selected periods[1]
(Percent per year)

Measure	1948-90	1948-73	1973-79	1979-90
Output per hour	2.2	3.1	0.6	0.9
Contribution of:				
Labor composition ...	.2	.2	.0	.4
Capital intensity	.9	1.1	.6	.5
Multifactor productivity	1.1	1.8	.1	.0

[1]Excludes government enterprises.

NOTE: Contributions may not sum to output per hour due to rounding.

[6] The notion of separability of inputs in a production function is developed in Ernst Berndt and Laurits Christensen, "The Internal Structure of Functional Relationships: Separability, Substitution, and Aggregation," *Review of Economic Studies*, Vol. 40, No. 3 (July 1973), pp. 403-410.

[7] Tables 23, 26, and 27 show multifactor productivity (MFP) measures net of the effect of labor composition. MFP measures are periodically updated, and in the future, MFP measures will be presented either gross or net of labor composition change depending on the purpose of future presentations.

Increases in the index of labor composition reflect changes in the educational attainment and level of work experience of the work force. Even though an exact decomposition of the separate contributions of education and experience is not possible, it appears likely that labor productivity growth was aided by increases in educational attainment while being hindered by declining levels of estimated work experience.[8] In the private business sector, the mean level of educational attainment for men rose from 9.3 years in 1948 to 13.0 years in 1990. Similarly, educational attainment of women increased from 9.9 to 13.0 years over the same period. The average number of years of estimated work experience for men fell from 18.8 years in 1948 to 17.8 years in 1990, while the average work experience for women fell slightly from 12.4 to 12.1 years. For all workers, the average level of estimated work experience declined from 17.4 years in 1948 to 15.6 years in 1990.

Capital intensity, or capital services per hour at work, increased at the annual rate of 2.7 percent for the period 1948-90. When weighted by capital's share of cost (about a third), the contribution of capital intensity added 0.9 percentage point to the growth rate of output per hour for the entire period. As a result, changes in capital intensity account for about 40 percent of the growth of labor productivity in the private business sector.

Multifactor productivity is the largest single source of labor productivity growth. Multifactor productivity increased 1.1 percent annually during the postwar period and contributed the same amount to output per hour. The increase in multifactor productivity accounts for half of the increase in labor productivity. Tables 26 and 27 present the index and percentage change for multifactor productivity as well as indexes for capital services, labor services, and labor composition for private business.

Changes in the growth rate of labor productivity

As shown in chart 9, the 2.2-percent annual average increase in private business labor productivity masked some sharp changes that occurred over the period. From 1948 to 1973, labor productivity increased 3.1 percent annually. Between 1973 and 1979, labor productivity grew at an annual rate of only 0.8 percent. In the most recent period, from 1979 to 1990, the growth rate of labor productivity continued at a sluggish pace of 0.9 percent per year.

In addition, labor composition can explain only a small portion of the slowdown. Before 1973, increases in workers' skills contributed 0.2 percentage point to the labor productivity growth rate, but during the 1973-79 period, the growth of labor composition stalled and did not contribute to productivity growth. The 0.2 percentage point drop in labor composition growth explains 9 percent of the 2.2-percentage-point slowdown in labor productivity growth (chart 10). For the most recent period 1979-90, labor composition has added 0.4 percentage point to the growth rate of output per hour, and it is a source of faster rather than slower labor productivity growth.

Changes in capital intensity also contributed to the slowdown in productivity growth. Prior to 1973, capital intensity contributed 1.1 percent annually to the growth rate of output per hour in the private business sector. During the 1973-79 period, capital intensity added only 0.6 percent. The 0.5-percentage-point decline explains less than 20 percent of the productivity slowdown. After 1979, the contribution of capital intensity declined further to 0.5 percentage point.

The rate of growth of multifactor productivity declined sharply after 1973, and it is the most important source of the labor productivity slowdown. Multifactor productivity increased 1.8 percent annually before 1973, but increased only 0.1 percent annually from 1973 to 1979. The 1.7-percentage-point slowdown in the annual growth rate accounts for more than three-quarters of the labor productivity slowdown during the 1973-79 period. During the most recent period, multifactor productivity did not increase and continues to be the single most important source of continuing slow labor productivity growth.

Sources of labor productivity growth in private nonfarm business

Table 24 shows essentially the same pattern for productivity trends in the private nonfarm business sector as seen in the business sector. Labor productivity in nonfarm business during the last 42 years grew at a slightly slower pace of 1.8 percent annually. Labor composition added 0.2 percentage point to the growth rate of output per hour and accounted for about 10 percent of the labor productivity increase. Capital intensity grew 2.5 percent annually and contributed 0.8 percentage point to the labor productivity growth rate or more than 40 percent of the growth in labor productivity. Multifactor productivity growth contributed as much as capital intensity, adding 0.8 percentage point to the growth rate of labor productivity or more than 40 percent of the increase in output per hour.

The growth rate of labor productivity slowed from 2.5 percent prior to 1973 to just 0.6 percent in the 1973-79 period in

Table 24. Sources of labor productivity growth in private nonfarm business, 1948-90 and selected periods[1]
(Percent per year)

Measure	1948-90	1948-73	1973-79	1979-90
Output per hour	1.8	2.5	0.6	0.7
Contribution of:				
Labor composition	.2	.2	.0	.4
Capital intensity	.8	.9	.6	.5
Multifactor productivity	.8	1.5	.0	-.1

[1] Excludes government enterprises.

NOTE: Contributions may not sum to output per hour due to rounding.

[8] See appendix H for a discussion of the limitations of decomposing labor composition growth into separate contributions of education and experience.

Chart 9. Sources of labor productivity growth in private business, 1948-90

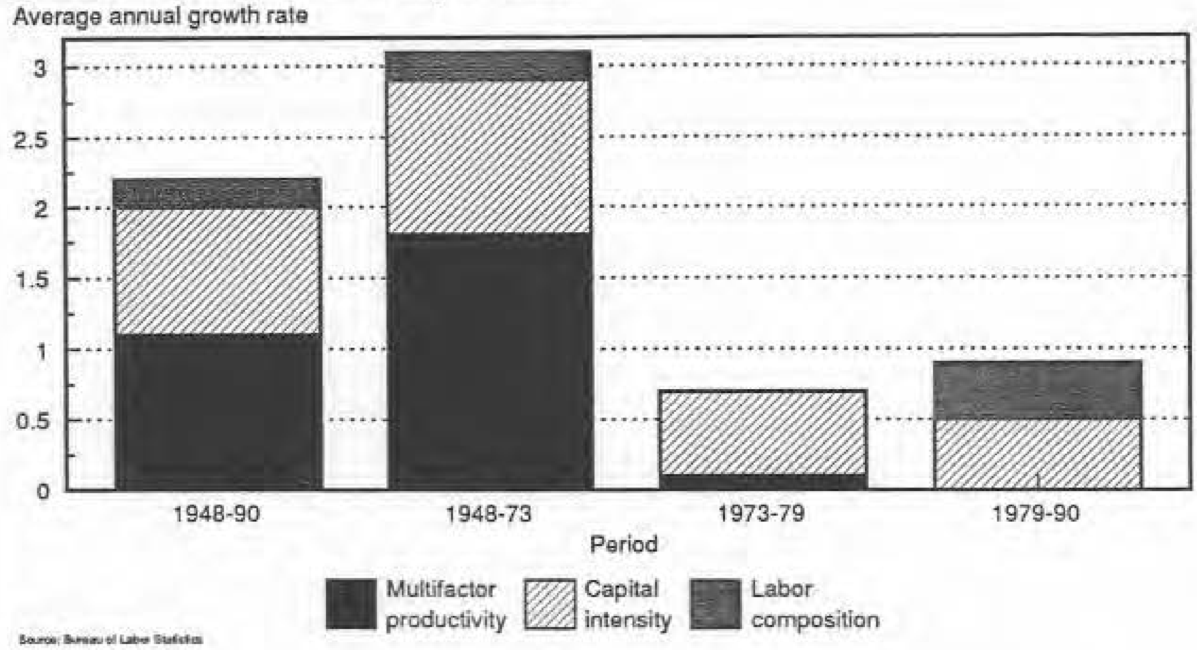
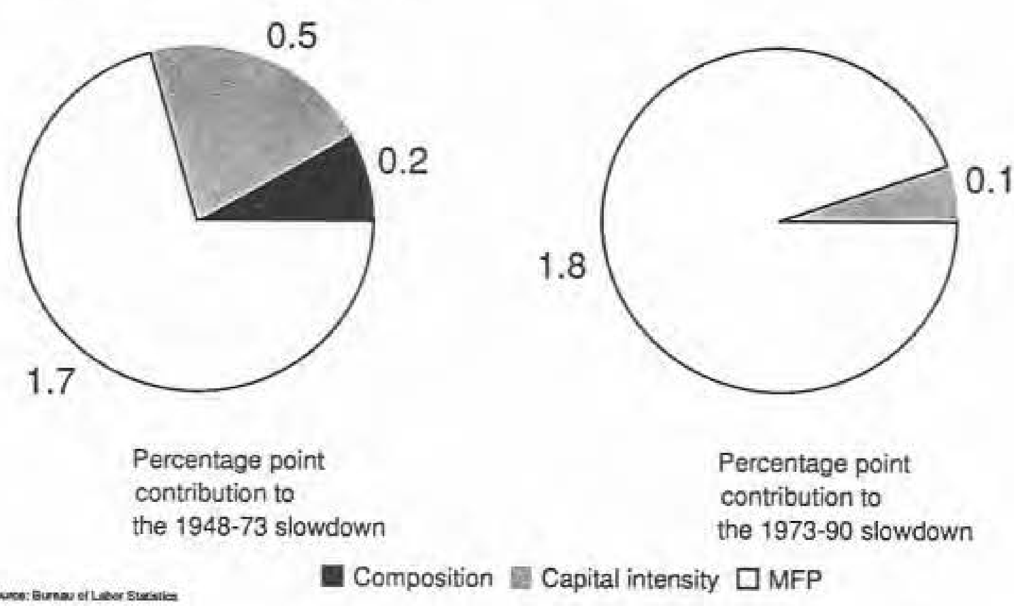


Chart 10. Contribution to changes in labor productivity growth rates in private business, 1948-90



private nonfarm business (chart 11). Labor composition accounted for 0.2 percentage point or about 10 percent of the 1.9-percent slowdown in output per hour (chart 12). After increasing at an annual rate of 1.5 percent prior to 1973, multifactor productivity did not change between 1973 and 1979. The 1.5-percentage-point decline in the growth rate of multifactor productivity accounted for the vast majority, more than 75 percent, of the slowdown in output per hour. Capital intensity contributed 0.3 percentage point or 15 percent to the slower growth.

Since 1979, the labor productivity growth rate in private nonfarm business has increased slowly, 0.7 percent per year. Labor composition grew at its fastest rate of the postwar period but still contributed only 0.4 percentage point annually. Capital intensity added 0.5 percentage point to the growth rate of output per hour, while multifactor productivity declined 0.1 percentage point. As a result of these changes, changes in labor composition were more than able to offset further declines in the contributions of multifactor productivity and capital intensity. Tables 28 and 29 provide indexes and percentage changes for multifactor productivity measures and factor inputs for private nonfarm business.

Relationship between factor prices and output per hour

Changes in the cost of labor can influence the demand for labor. In a competitive economy, firms act as price-takers and attempt to minimize the cost of production at the prevailing factor prices. Relative factor prices, in particular the cost of capital relative to the cost of labor, can change due to shifts in the supply of labor and capital or nonneutral technical change. Cost minimization implies a relationship between factor prices and their relative demand. The upper half of table 25 shows that adjusted hourly compensation (average hourly compensation divided by the index of labor composition) has consistently grown faster than the price of capital services. For the entire period, adjusted hourly compensation grew at an annual rate of 6.0 percent, while the price of capital services increased 2.9 percent annually. The relative price of capital to labor declined 3.1 percent annually. The lower half of table 25 shows that firms have increased their use of capital services at a faster pace than labor input. Annual increases in capital services averaged 3.7 percent, while labor grew only 1.3 percent. Capital intensity increased at an annual rate of 2.4 percent.

Increases in capital intensity have slowed concurrently with increasingly smaller declines in the relative factor prices. Relative factor prices declined 4.0 percent annually before 1973, but during the next 6 years relative factor prices declined at the rate of only 2.6 percent. Since 1979 the relative price of capital to labor has continued to decline but at an even slower rate of 1.3 percent. Similarly, capital intensity, which increased 3.0 percent annually prior to 1973, rose 2.0 percent annually from 1973 to 1979. After 1979, the growth rate of capital intensity slowed further to 1.3 percent.

Table 25. Relationship between changes in factor prices and changes in factor inputs in private business, 1948-90 and selected periods
(Percent per year)

Measure	1948-90 (1)	1948-73 (2)	1973-79 (3)	1979-90 (4)	Change 1948-73 to 1973-79 (3)-(2)	Change 1973-79 to 1979-90 (4)-(3)
Factor prices						
Capital services[1]	2.9	1.7	6.5	3.8	4.8	-2.7
Labor services[2]	6.0	5.7	9.1	5.1	3.4	-4.0
Ratio of price of capital to price of labor[3]	-3.1	-4.0	-2.6	-1.3	-1.4	-1.3
Factor inputs						
Capital services	3.7	3.9	3.8	3.3	-.1	-.5
Labor services[4]	1.3	.9	1.8	2.0	.9	.2
Ratio of capital to labor services	2.4	3.0	2.0	1.3	-1.0	-.7

[1] Implicit price of capital services in private business defined as capital compensation divided by the Tornqvist index of capital services.
[2] Hourly compensation of all persons adjusted for changing composition of the private business work force.
[3] The ratio of the implicit price of capital services to adjusted hourly compensation.
[4] Tornqvist index of labor input.

Chart 11. Sources of labor productivity growth in private nonfarm business, 1948-90

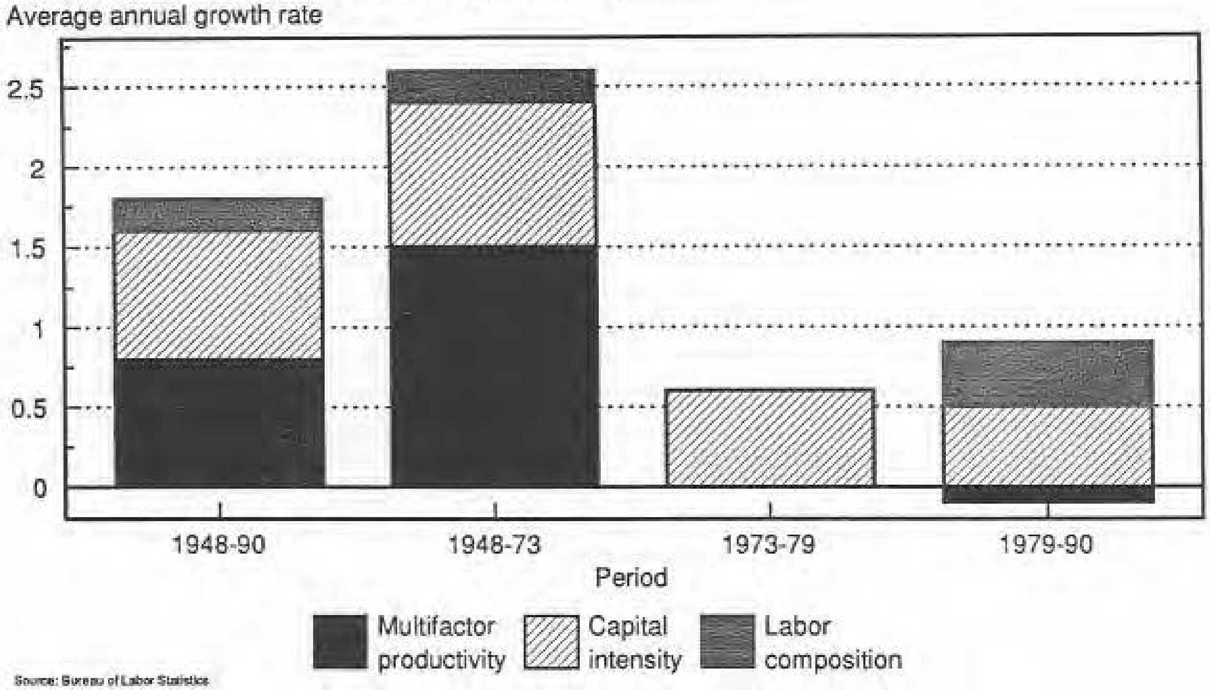


Chart 12. Contribution to changes in labor productivity growth rates in private nonfarm business, 1948-90

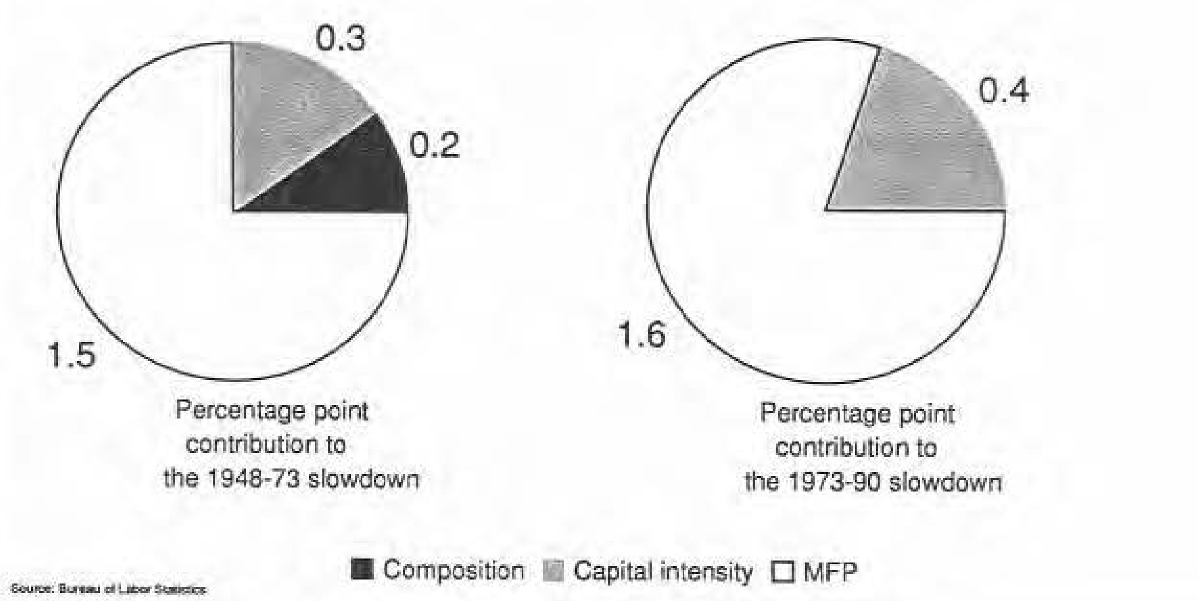


Table 26. Private business sector: Productivity and related measures, 1948-90[1]
Indexes 1982=100

Year	Productivity		Output[3]	Factor inputs			Labor	
	Output per hour of all persons	Multifactor productivity[2]		Labor input[4]	Capital services[5]	Combined units of labor and capital inputs[6]	Hours of all persons[7]	Composition[8]
1948	44.5	66.7	35.6	73.0	27.0	53.4	80.0	91.3
1949	45.2	66.0	34.9	71.2	27.6	52.9	77.8	91.5
1950	49.1	70.7	38.3	71.9	28.9	54.2	78.1	92.2
1951	50.9	72.5	40.9	74.1	30.7	56.4	80.2	92.4
1952	52.9	74.1	42.4	74.2	31.8	57.1	80.1	92.7
1953	54.6	75.6	44.1	75.4	32.8	58.3	80.8	93.4
1954	56.0	75.8	43.7	73.4	33.5	57.7	78.1	94.0
1955	57.9	78.3	46.9	76.3	34.8	59.9	81.0	94.2
1956	58.7	78.6	48.2	77.6	36.3	61.4	82.2	94.4
1957	60.4	79.4	48.9	76.6	37.6	61.5	80.8	94.8
1958	62.4	80.0	48.0	73.3	38.3	60.1	77.0	95.2
1959	64.0	82.3	51.2	76.2	39.3	62.2	80.0	95.3
1960	65.1	82.6	52.0	76.6	40.5	63.0	80.0	95.8
1961	67.6	84.4	53.2	75.8	41.5	63.0	78.7	96.3
1962	70.0	86.3	55.9	77.7	42.8	64.7	79.9	97.2
1963	72.9	88.9	58.5	78.2	44.3	65.8	80.3	97.4
1964	76.1	92.2	62.0	79.4	46.1	67.3	81.5	97.5
1965	78.2	94.2	65.7	81.9	48.3	69.8	84.1	97.4
1966	80.3	95.7	69.1	83.7	51.2	72.2	86.0	97.3
1967	82.4	96.3	70.7	83.6	54.1	73.4	85.7	97.5
1968	85.0	98.5	73.9	84.5	56.7	75.0	86.8	97.3
1969	85.5	98.0	76.1	86.9	59.4	77.6	89.0	97.7
1970	86.8	97.3	75.7	85.5	62.1	77.7	87.2	98.1
1971	89.7	99.6	77.9	84.9	64.4	78.2	86.8	97.8
1972	92.5	102.4	82.9	87.8	66.9	81.0	89.7	97.8
1973	94.9	104.8	88.2	90.7	70.4	84.1	92.9	97.6
1974	93.2	101.0	86.6	91.3	73.9	85.7	92.9	98.3
1975	95.4	101.0	84.9	87.4	76.4	84.0	88.9	98.3
1976	99.2	104.2	89.9	89.8	78.5	86.3	91.5	98.1
1977	99.8	106.1	94.9	93.3	81.1	89.5	95.1	98.1
1978	100.4	106.9	100.1	98.0	84.3	93.7	99.7	98.2
1979	99.3	105.6	102.2	100.8	88.0	96.8	102.9	97.9
1980	98.5	102.7	100.4	100.1	92.5	97.8	101.9	98.2
1981	99.8	102.3	102.4	101.6	96.6	100.0	102.6	99.0
1982	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0
1983	102.3	101.8	104.2	102.2	102.4	102.3	101.8	100.4
1984	104.7	104.8	112.8	108.3	106.0	107.6	107.7	100.5
1985	106.3	105.7	116.8	110.8	110.1	110.6	109.9	100.8
1986	108.5	106.5	120.1	112.0	114.2	112.7	110.6	101.3
1987	109.6	107.5	125.0	115.8	117.4	116.3	114.0	101.6
1988	110.6	108.1	130.4	120.7	120.4	120.6	117.9	102.4
1989	109.7	106.9	132.6	124.4	123.3	124.0	120.9	102.9
1990	110.0	106.1	133.0	125.0	126.0	125.3	120.9	103.4

See footnotes following table 29.
Source: Bureau of Labor Statistics

Table 27. Private business sector: Productivity and related measures, 1949-90[1]
(Percent change)

Year	Productivity		Output[3]	Factor inputs			Labor	
	Output per hour of all persons	Multifactor productivity[2]		Labor input[4]	Capital services[5]	Combined units of labor and capital inputs[6]	Hours of all persons[7]	Composition[8]
1949	1.6	-1.1	-1.9	-3.2	2.4	-0.8	-3.4	0.2
1950	8.6	7.2	9.7	1.8	4.7	2.3	1.0	.8
1951	3.7	2.4	6.6	3.1	6.0	4.1	2.8	.3
1952	3.8	2.3	3.7	.1	3.7	1.3	-.1	.3
1953	3.2	2.0	4.1	1.6	3.1	2.1	.9	.8
1954	2.6	.3	-.9	-2.6	2.3	-1.1	-3.4	.7
1955	3.4	3.2	7.2	3.9	3.8	3.9	3.7	.2
1956	1.3	.3	2.9	1.6	4.3	2.5	1.5	.2
1957	2.9	1.1	1.3	-1.2	3.5	.2	-1.6	.4
1958	3.2	.7	-1.7	-4.4	2.1	-2.4	-4.8	.4
1959	2.7	2.9	6.6	3.9	2.6	3.6	3.8	.1
1960	1.6	.3	1.7	.6	3.1	1.3	.0	.5
1961	3.9	2.2	2.1	-1.2	2.3	.0	-1.7	.6
1962	3.5	2.4	5.2	2.5	3.3	2.7	1.6	.9
1963	4.2	3.0	4.7	.7	3.5	1.6	.5	.2
1964	4.4	3.6	6.0	1.6	3.9	2.3	1.5	.0
1965	2.7	2.2	6.0	3.0	5.0	3.7	3.1	-.1
1966	2.8	1.6	5.2	2.2	6.0	3.4	2.3	-.0
1967	2.6	.6	2.2	-.2	5.6	1.6	-.3	.2
1968	3.1	2.2	4.5	1.1	4.7	2.3	1.4	-.2
1969	.6	-.5	3.0	2.8	4.9	3.4	2.4	.4
1970	1.5	-.7	-.5	-1.5	4.4	.2	-1.9	.4
1971	3.4	2.3	2.9	-.8	3.8	.6	-.5	-.3
1972	3.1	2.9	6.5	3.3	3.9	3.6	3.3	.0
1973	2.6	2.4	6.3	3.4	5.1	3.9	3.6	-.2
1974	-1.8	-3.7	-1.8	.7	5.0	2.0	.0	.7
1975	2.4	.0	-2.0	-4.2	3.4	-2.0	-4.2	.0
1976	2.9	3.1	5.9	2.6	2.7	2.7	2.9	-.2
1977	1.6	1.8	5.6	3.9	3.3	3.7	3.9	.0
1978	.5	.8	5.5	5.0	4.0	4.7	4.9	.1
1979	-1.1	-1.2	2.0	2.8	4.3	3.3	3.2	-.3
1980	-.7	-2.7	-1.7	-.6	5.1	1.0	-1.0	.3
1981	1.3	-.4	2.0	1.5	4.5	2.3	.7	.8
1982	.2	-2.3	-2.3	-1.5	3.5	.0	-2.5	1.0
1983	2.3	1.8	4.2	2.2	2.4	2.3	1.8	.4
1984	2.4	2.9	8.2	5.9	3.5	5.2	5.8	.1
1985	1.5	.8	3.6	2.4	3.8	2.8	2.1	.2
1986	2.1	.8	2.8	1.2	3.7	1.9	.7	.5
1987	1.0	.9	4.1	3.3	2.8	3.2	3.1	.2
1988	.9	.6	4.3	3.8	2.5	3.7	3.0	.8
1989	-.8	-1.1	1.7	3.0	2.5	2.9	2.5	.5
1990	.3	-.8	.3	1.1	2.2	1.0	.6	.5

See footnotes following table 29.
Source: Bureau of Labor Statistics

Table 28. Private nonfarm business sector: Productivity and related measures, 1948-90[1]
Indexes 1982=100

Year	Productivity		Output[3]	Factor inputs			Labor	
	Output per hour of all persons	Multifactor productivity[2]		Labor input[4]	Capital services[5]	Combined units of labor and capital inputs[6]	Hours of all persons[7]	Composition[8]
1948	51.6	73.8	35.0	62.6	25.2	47.4	67.8	92.3
1949	52.8	73.8	34.3	60.1	25.7	46.5	65.1	92.4
1950	56.2	78.0	37.7	62.3	26.9	48.4	67.1	92.8
1951	57.6	79.4	40.4	65.2	28.6	50.9	70.2	92.8
1952	59.3	80.6	41.9	66.3	29.7	52.0	70.7	93.5
1953	60.2	81.2	43.6	68.3	30.6	53.7	72.5	94.3
1954	61.8	81.5	43.2	66.4	31.4	53.0	69.9	94.9
1955	63.8	84.0	46.4	69.2	32.6	55.3	72.8	95.1
1956	64.1	83.8	47.8	71.1	34.1	57.1	74.6	95.3
1957	65.4	84.1	48.5	71.0	35.4	57.7	74.1	95.8
1958	67.3	84.6	47.7	68.0	36.2	56.4	70.9	95.9
1959	68.8	86.9	50.9	71.0	37.2	58.6	74.0	96.0
1960	69.6	87.0	51.7	71.4	38.4	59.4	74.4	96.0
1961	71.9	88.4	52.8	71.2	39.4	59.8	73.5	96.9
1962	74.2	90.2	55.6	73.4	40.8	61.7	75.0	97.9
1963	76.9	92.6	58.3	74.4	42.3	63.0	75.8	98.1
1964	79.8	95.5	61.9	76.1	44.0	64.8	77.6	98.0
1965	81.6	97.3	65.6	78.8	46.3	67.5	80.5	97.9
1966	83.1	98.4	69.2	81.3	49.2	70.3	83.2	97.7
1967	84.9	98.8	70.6	81.3	52.1	71.5	83.2	97.7
1968	87.5	101.0	74.0	82.3	54.7	73.2	84.5	97.4
1969	87.5	100.1	76.1	85.1	57.5	76.1	87.0	97.8
1970	88.5	99.1	75.7	84.0	60.2	76.4	85.5	98.2
1971	91.3	101.2	77.9	83.5	62.6	77.0	85.3	97.9
1972	94.0	104.1	83.0	86.4	65.2	79.7	88.2	97.9
1973	96.4	106.4	88.4	89.6	68.7	83.1	91.7	97.7
1974	94.5	102.3	86.7	90.2	72.4	84.7	91.7	98.4
1975	96.7	102.3	84.9	86.4	75.0	83.0	87.8	98.4
1976	99.2	105.3	90.0	89.0	77.2	85.5	90.7	98.1
1977	100.6	106.9	95.1	92.7	80.0	88.9	94.5	98.1
1978	101.3	107.9	100.6	97.5	83.4	93.2	99.3	98.2
1979	99.8	106.1	102.6	100.5	87.6	96.7	102.7	97.9
1980	99.0	103.2	100.8	100.0	92.0	97.6	101.8	98.2
1981	99.8	102.4	102.4	101.5	96.3	100.0	102.6	98.9
1982	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0
1983	102.5	102.0	104.5	102.4	102.6	102.4	102.0	100.4
1984	104.6	104.8	113.2	108.7	106.4	108.0	108.2	100.4
1985	105.5	105.0	117.0	111.6	111.0	111.4	110.8	100.7
1986	107.6	105.7	120.2	113.1	115.2	113.7	111.7	101.3
1987	108.5	106.5	125.2	117.1	118.7	117.6	115.4	101.5
1988	109.4	107.0	130.8	122.2	122.0	122.2	119.5	102.3
1989	108.3	105.6	133.0	126.2	125.4	125.9	122.8	102.8
1990	108.3	104.5	133.1	127.0	128.3	127.4	122.9	103.3

See footnotes following table 29.
Source: Bureau of Labor Statistics

Table 29. Private nonfarm business sector: Productivity and related measures, 1949-90[1]
(Percent change)

Year	Productivity		Output[3]	Factor inputs			Labor	
	Output per hour of all persons	Multifactor productivity[2]		Labor input[4]	Capital services[5]	Combined units of labor and capital inputs[6]	Hours of all persons[7]	Composition[8]
1949	2.3	0.0	-1.8	-3.8	2.1	-1.8	-4.0	0.1
1950	6.6	5.7	9.8	3.6	4.7	4.0	3.1	.5
1951	2.4	1.8	7.1	4.6	6.2	5.2	4.6	.0
1952	2.9	1.5	3.8	1.5	3.8	2.2	.8	.7
1953	1.6	.7	4.1	3.3	3.3	3.3	2.5	.8
1954	2.6	.4	-1.0	-2.9	2.3	-1.3	-3.5	.7
1955	3.2	3.0	7.4	4.4	4.1	4.3	4.1	.2
1956	.5	-.2	3.0	2.6	4.6	3.3	2.5	.2
1957	2.1	.4	1.4	-.1	3.7	1.1	-.6	.5
1958	2.8	.6	-1.7	-4.2	2.2	-2.3	-4.4	.1
1959	2.3	2.7	6.8	4.5	2.8	3.9	4.4	.1
1960	1.1	.2	1.6	.6	3.3	1.4	.5	.0
1961	3.4	-1.5	2.1	-.3	2.5	.6	-1.2	1.0
1962	3.1	2.0	5.3	3.1	3.5	3.2	2.1	1.0
1963	3.7	2.7	4.7	1.2	3.7	2.0	1.0	.2
1964	3.8	3.2	6.2	2.3	4.0	2.9	2.3	.0
1965	2.3	1.9	6.1	3.6	5.2	4.1	3.7	-.1
1966	1.8	1.1	5.3	3.2	6.3	4.2	3.4	-.2
1967	2.2	.4	2.2	.0	5.9	1.8	-.1	.0
1968	3.0	2.3	4.7	1.3	4.9	2.4	1.6	-.3
1969	.0	-1.0	2.9	3.4	5.2	3.9	3.0	.4
1970	1.1	-1.0	-.6	-1.3	4.7	.4	-1.7	.4
1971	3.2	2.1	2.9	-.5	3.9	.7	-.3	-.3
1972	3.0	2.9	6.6	3.4	4.1	3.6	3.5	.0
1973	2.5	2.2	6.5	3.8	5.4	4.2	3.9	-.2
1974	-1.9	-3.8	-1.9	.7	5.4	2.0	.0	.7
1975	2.3	.0	-2.1	-4.3	3.6	-2.1	-4.3	.0
1976	2.6	2.9	6.0	3.0	2.9	3.0	3.3	-.3
1977	1.4	1.6	5.7	4.2	3.5	4.0	4.2	.0
1978	.7	.9	5.8	5.1	4.3	4.9	5.1	.1
1979	-1.4	-1.6	2.0	3.1	5.0	3.7	3.5	-.3
1980	-.9	-2.7	-1.8	.6	5.0	1.0	-.9	.3
1981	.9	-.8	1.6	1.6	4.7	2.4	.7	.8
1982	.2	-2.4	-2.4	-1.5	3.8	.0	-2.5	1.1
1983	2.5	2.0	4.5	2.4	2.6	2.4	2.0	.4
1984	2.1	2.7	8.3	6.1	3.8	5.4	6.1	.1
1985	.9	.2	3.4	2.7	4.3	3.2	2.5	.2
1986	2.0	.7	2.8	1.4	3.8	2.1	.8	.6
1987	.8	.7	4.1	3.5	3.0	3.4	3.3	.2
1988	.9	.5	4.5	4.4	2.8	3.9	3.5	.8
1989	-1.0	-1.4	1.7	3.3	2.7	3.1	2.7	.5
1990	.0	-1.0	.1	.6	2.3	1.1	.1	.5

Footnotes, tables 26-29.

Source: Output data are from the Bureau of Economic Analysis (BEA), U.S. Department of Commerce, and the Federal Reserve Board. Compensation and hours data are from the Bureau of Labor Statistics, U.S. Department of Labor, and BEA. Capital measures are based on data supplied by BEA and the U.S. Department of Agriculture.

[1] The private business sector includes all of Gross National Product except the output of general government, government enterprises, nonprofit institutions, the rest-of-the-world sector, the rental value of owner-occupied real estate, the output of paid employees of private households, and the statistical discrepancy in computing the national income and product accounts. The private nonfarm business sector also excludes farms but includes agricultural services.

[2] Output per unit of combined labor and capital inputs.

3 Gross Domestic Product originating in the sector, in constant dollars.

[4] A labor compensation share weighted average of all hours classified by education, work experience, and sex.

[5] A measure of the flow of capital services used in the sector.

[6] Hours at work of all persons combined with capital input, using labor's and capital's shares of costs as weights.

[7] Hours at work of all persons including employees, proprietors, and unpaid family workers engaged in the sector.

[8] A measure of the contribution to labor input of changes in the distribution of workers by education, work experience, and sex.

Summary

In chapter II, removing the assumption that workers are homogeneous led to a measure of weighted labor input which is the sum of total hours and labor composition. In this chapter, a growth accounting framework has been introduced which demonstrates that the contribution of labor input to output growth can be measured as the product of the growth rate of weighted labor input and labor's share of total costs. This contribution can be divided into two components: the contribution of total hours and the contribution of changing education, work experience, and sex composition of the work force. Labor composition contributes equally to both output and labor productivity.

Output per hour grew at an annual rate of 2.2 percent in the private business sector and 1.8 percent in the private nonfarm sector between 1948 and 1990. Charts 9 and 11 summarize the sources of labor productivity growth.[9] Changes in labor composition contributed about one-tenth of the growth in labor productivity. About half of the growth of productivity is accounted for by increases in multifactor productivity. The growth of capital intensity accounted for the remainder.

The rate of labor productivity growth slowed after 1973. Most of this slowdown can be attributed to slower multifactor productivity growth which has stagnated since 1973. The growth rate of labor composition declined during the 1973-79 period and so accounted for approximately one-tenth of the slowdown. Since 1979, the growth rate of labor productivity has continued its weak performance. The slightly faster labor productivity growth is due to more rapid growth in labor composition which more than offset further declines in the contributions of multifactor productivity and capital intensity.

[9] In this bulletin, multifactor productivity, capital, and labor composition are identified as three sources of labor productivity growth. A more detailed decomposition of labor productivity growth might have revealed other important sources of labor productivity growth.

Appendix A: The Labor Composition Model

The measurement of labor composition is part of a continuing Bureau of Labor Statistics (BLS) effort to identify and measure sources of labor productivity growth. Labor composition extends the measurement of labor input and productivity change because labor input is now based on a more detailed model of production than previously used by the BLS (see *Trends in Multifactor Productivity, 1948-81*). Rather than simply identifying labor input with hours, labor services are permitted to vary across workers. The measurement of labor input and labor composition effectively requires workers to be partitioned into distinct groups and measurement of the hours and associated costs for each type of worker.

The measures developed in this bulletin show how changes in the composition of the work force affect labor input and, ultimately, productivity. The composition of the work force has changed in many ways, but this study focuses solely on changes in the educational attainment and work experience of both men and women. While other characteristics of the work force could have been added to the study and resulted in a broader measure of labor composition change, the sources of labor composition change would have been less clear. Accordingly, this study has narrowed its focus so that the impact of rising educational attainment, the baby boom and subsequent baby bust, and rising female labor force participation rates on measures of labor input and productivity can be examined.

This appendix begins with a brief review of the growth accounting framework and the derivation of the Tornqvist index of labor input. The remainder is devoted to a discussion of the disaggregation of hours focusing primarily on determining the set of characteristics that is most appropriate for partitioning workers.

Labor input index

A description of the production process can be summarized by a mathematical statement known as a production function. The production function describes how much output specific quantities of capital and labor can generate. In previous BLS multifactor productivity work (Bulletin 2178, cited above), the production function was assumed to be:

$$(1)\ q = f(k_1, \ldots, k_n, H, t)$$

where output q is generated by n types of capital input, $k_1, \ldots, k_n$, a single type of labor input, H, and the technology available at time t.

In the production function described by equation 1, output stays the same regardless of who performs the work. For example, replacing a high school graduate with a college graduate or a young worker with an experienced worker would leave output unchanged. In effect, the production function treats labor as a homogeneous input and ignores productive differences among workers.

The labor composition model is formed by generalizing equation 1 to permit numerous types of labor input. This can be written as:

$$(2)\ q = g(k_1, \ldots, k_n, h_1, \ldots, h_m, t)$$

where output q is generated by n types of capital inputs, $k_1, \ldots, k_n$, m types of labor inputs, $h_1, \ldots, h_m$, and the technology available at time t. The production function in equation 2 allows each type of labor input to have a unique effect on output. Therefore, when differences in workers do significantly affect output, equation 2 represents the production function better than equation 1.

The growth accounting framework relates the growth rate of output to the growth rates of each input and productivity in a manner consistent with the most general description of equation 2. By taking the logarithm and then the derivative with respect to time, equation 2 can be expressed in terms of growth rates:

$$(3)\ \dot{q}/q = \partial g/\partial k_1 * \dot{k}_1/k_1 + \ldots + \partial g/\partial k_n * \dot{k}_n/k_n + \partial g/\partial h_1 * \dot{h}_1/h_1 + \ldots + \partial g/\partial h_m * \dot{h}_m/h_m + \dot{A}/A$$

For any input x_1, the dot notation such as $\dot{x}_1/x_1$ denotes the growth rate of the variable x_1. The term $\partial g/\partial h_1$ refers to the partial derivative of the logarithm of the production function with respect to an input such as h_1. This is also known as the output elasticity, which measures the percentage change in output for a 1-percent change in an input. The effect of multifactor productivity on output is measured by the partial derivative of the log of the production function with respect to time, $\partial g/\partial t * 1/q$. Rather than using this cumbersome notation, a common convention is to express the growth rate of multifactor productivity using the shorthand term, $\dot{A}/A$.

Rearranging equation 3 yields an expression for the growth rate of multifactor productivity.

$$(4)\ \dot{A}/A = \dot{q}/q - \partial g/\partial k_1 * \dot{k}_1/k_1 - \ldots - \partial g/\partial k_n * \dot{k}_n/k_n - \partial g/\partial h_1 * \dot{h}_1/h_1 - \ldots - \partial g/\partial h_m * \dot{h}_m/h_m$$

The output elasticities cannot be directly observed. Instead, it is assumed that the production function exhibits constant returns to scale and that factor input markets are in competitive equilibrium. By further assuming cost minimizing behavior, one can demonstrate that each factor's output elasticity will equal its share of total costs. Designating s_{ki} and s_{hj} as factor shares for capital and labor inputs, respectively, equation 4 can be written as:

$$(5)\ \dot{A}/A = \dot{q}/q - s_{k1} * \dot{k}_1/k_1 - \ldots - s_{kn} * \dot{k}_n/k_n - s_{h1} * \dot{h}_1/h_1 + \ldots - s_{hm} * \dot{h}_m/h_m$$

where the factor costs shares are:

$$s_{ki} = P_{ki} * k_i / \Sigma_{ij} (P_{ki} * k_i + P_{hj} * h_j)$$
$$s_{hj} = P_{hj} * h_j / \Sigma_{ij} (P_{ki} * k_i + P_{hj} * h_j)$$

and P_{ki} and P_{hj} are the prices of capital and labor services for the i^{th} asset and the j^{th} type of labor, and k_i and h_j are quantities of those capital and labor services.

Assuming separability of inputs and Hicks neutral technical change permits meaningful and unambiguous aggregates of capital and labor input. That is, aggregates of labor and capital can be written as:

$$(6a)\ \dot{L}/L = s_{l1} * \dot{h}_1/h_1 + \ldots + s_{lm} * \dot{h}_m/h_m$$
$$(6b)\ \dot{K}/K = s_{c1} * \dot{k}_1/k_1 + \ldots + s_{cn} * \dot{k}_n/k_n$$

where s_{li} is the share of *labor compensation* paid to the i^{th} labor input rather than the share of *total cost* used in equation 5, and similarly, s_{cj} is the capital costs share rather than total cost share. Equation 5 can be now written more simply.

$$(7)\ \dot{A}/A = \dot{Q}/Q - s_k * \dot{K}/K - s_l * \dot{L}/L$$

where s_l and s_k are labor's and capital's shares of total costs.

To implement equation 7, a specific form of the production function (equation 2) must be chosen. The translog function imposes less restrictive assumptions than other common functional forms such as the Cobb-Douglas or constant elasticity of substitution (CES) production functions.[1] Diewert has shown that certain index number formulas are consistent with particular functional forms of the production function. In particular, the Tornqvist index, which is consistent with the translog production function, belongs to a class of so-called "superlative" index number formulas.[2] That is, changes in output consistent with the very general translog production function are exactly measured by changes in Tornqvist indexes. Other index number formulas such as fixed weight indexes have been shown to be consistent with more restrictive production functions. Use of fixed weight index number formulas such as the Paasche or Lespeyres can result in an index of an aggregate which contains a very substantial "index number bias," especially during periods of rapid change in input prices or quantities. Accordingly, changes in multifactor productivity as well as aggregate labor and capital inputs are measured with Tornqvist indexes. Instantaneous growth rates, such as $\dot{L}/L$, must be replaced by annual rates of change. In the Tornqvist index number formula, these are measured as differences in successive natural logarithms, that is, $\Delta \ln L = \ln L_t - \ln L_{t-1}$. Equations 6a and b are rewritten as:

$$(8a)\ \Delta \ln L = \Sigma_j\ 1/2 * (s_{lj}(t) + s_{lj}(t\text{-}1))\ \Delta \ln h_j$$
$$(8b)\ \Delta \ln K = \Sigma_i\ 1/2 * (s_{ci}(t) + s_{ci}(t\text{-}1))\ \Delta \ln k_i$$

and the Tornqvist index of multifactor productivity growth is:

$$(9)\ \Delta \ln A = \Delta \ln Q - 1/2 * (s_k(t) + s_k(t\text{-}1)) * \Delta \ln K - 1/2 * (s_l(t) + s_l(t\text{-}1)) * \Delta \ln L$$

Changes in the index of labor composition, LC, are defined as the difference between changes in the aggregate labor input index, L, and the simple sum of the hours of all persons, H.

$$(10)\ \Delta \ln LC = \Delta \ln L - \Delta \ln H \quad \text{or} \quad = \Delta \ln L/H$$

Labor composition has an important interpretation in terms of its role in explaining the differences between equations 1 and 2. BLS has heretofore measured multifactor productivity (MFP) growth ($\dot{B}/B$) by applying steps analogous to equations 3 through 9 to equation 1 rather than equation 2. Hence,

$$(11)\ \Delta \ln B = \Delta \ln Q - 1/2 * (s_k(t) + s_k(t\text{-}1)) * \Delta \ln K - 1/2 * (s_l(t) + s_l(t\text{-}1)) * \Delta \ln H$$

It follows by comparing equations 11 and 9 that:

$$(12)\ \Delta \ln B = \Delta \ln A + 1/2 * (s_l(t) + s_l(t\text{-}1)) * \Delta \ln LC$$

As Robert Solow (1958) pointed out, the term $\Delta \ln B$ measures shifts over time in the function f of equation 1. Similarly, $\Delta \ln A$ measures shifts in the function g in equation 2. Therefore, previous BLS MFP measures ($\Delta \ln B$) have two components: change due to shift in the "correctly" specified function g and change due to changes in labor composition.

In terms of this formal production theory, $\Delta \ln A$ is preferred because it corrects for a bias inherent in equation 1. Since Solow's original article, however, the purpose of MFP measurement has usually been to sort out the sources of output per hour growth. As Solow made clear, MFP will reflect the effects of many unspecified influences including returns to scale and capacity utilization, and will even reflect measurement errors in both outputs and inputs. Therefore, Solow stressed an interpretation of MFP measurement as a way of sorting out one of the sources of output per hour growth (capital). In light of that, the new MFP measure, $\dot{A}/A$, is that part of $\dot{B}/B$ which cannot be attributed to changes in labor composition.

[1] Christensen, Jorgenson, and Lau (1973) demonstrate the advantages of a translog production function.

[2] See Diewert (1976) for a derivation of the relationship between translog production functions and Tornqvist index numbers. See Caves, Christensen, and Diewert (1982) and Denny and Fuss (1983) for a discussion of some limitations and restrictions of the Tornqvist index.

Increases in labor composition result from labor input growing faster than total hours. The quantity of labor services per hour is increasing whenever $\Delta \ln L/H > 0$. An increase in the quantity of services per hour is consistent with an increase in the marginal product of labor, that is, an increase in the average of the marginal products of all workers.

As an example of this, imagine replacing one high school graduate with one college graduate. Because the college graduate presumably has a higher marginal product than the high school graduate (holding all else constant), the marginal product of the work force will rise. As a result, each hour of the higher paid college graduate is weighted more heavily than an hour of the lower paid high school graduate in the measures of labor input. Since the changes in hours of the two workers are symmetrical, weighting hours which are increasing more heavily than hours which are declining leads to an increase in labor input as can be seen in equation 8a. By assumption, total hours of all workers remain unchanged, and so equation 10 shows that an increase in aggregate labor input without any change in total hours will also produce an increase in labor composition. Labor composition change can thus be viewed as the change in the average amount of labor services associated with an hour of work.

Disaggregation of labor

For measuring labor input, a distinction could be made when two groups of workers differ in their marginal products. When labor markets are competitive, earnings differences reflect differences in marginal products. In theory then, any observed earnings difference between groups of labor inputs forms a basis for treating the groups separately. In practice, there are far too many sources of earnings differentials to use them all. In addition, some observed earnings differentials are incompatible with a perfectly competitive labor market.[3]

It is impractical to disaggregate labor by every (or even many) observable earnings difference, thus the measurement of labor composition involves making some choices as to which earnings differences to focus on. Workers may be grouped in categories differentiated by various characteristics. From the perspective of formal production theory, it might be thought desirable to group workers in such a way that the "most important" earnings differences are distinguished. Presumably that would yield a good approximation to the "true" labor composition measure. Jorgenson, Gollop, and Fraumeni (1987) have estimated labor composition using a very large number of categories representing a cross-classification of five characteristics (age, education, class of worker, occupation, and sex) for each industry. From the perspective of production theory, this has the advantage that labor composition will reflect not only the direct contribution of the specified characteristics, but also include the many effects of correlated traits not specifically included within the framework. For example, the contribution of changing gender distribution to labor composition may also incorporate shifts in the proportion of workers working part time.

For this study, a somewhat different approach is adopted, linked closely to the goal of explaining the sources of output per hour growth. A comprehensive set of traits can lead to difficulties in identifying sources of labor composition growth. In the previous example, shifts in the proportion of part-time employment may be a source of labor composition change even though work schedules are not a classifying characteristic. Instead, this study narrows the focus in order to identify and measure the effects on productivity of a specific known set of factors. Thus, rather than partition MFP growth ($\dot{B}/B$) into two catch-all categories ("labor composition effects" and "all other sources of MFP growth"), a labor composition measure is developed that attempts to identify the separate effects and interacted effects on productivity of changes in two specific characteristics: education and work experience. Because earnings of men and women differ markedly, the method is applied separately to each group.

Human capital model

Human capital theory is used in this study as an organizing principle for partitioning groups of workers. The human capital literature as well as the hedonic wage literature have repeatedly documented earnings differences by education and work experience.[4] Furthermore, the human capital model ties these earnings differentials to differences in marginal products of workers in competitive labor markets.

The work of Becker (1975) on the theory of human capital best codifies the role of education and on-the-job training in the acquisition of skills and earnings. In human capital theory, skills are the ultimate source of worker productivity. Education and training are the means for acquiring additional skills beyond innate abilities. Firms offer higher wages because the additional skills embodied in larger amounts of education and training raise the marginal product of these workers above those without such skills. Education and training are an investment in which current earnings are foregone in favor of higher future earnings. Workers presumably choose levels of education and training which maximize the present discounted value of their earnings. The optimal quantity of education and training for a worker depends on the wages offered by firms for each level of education and training, the worker's ability to acquire skills through education and training, the discount rate or market rate of interest, the expected length of his or her working life, and other factors. In long run competitive equilibrium, differences in earnings by education and training reflect differences in marginal products, and the

[3] For example, imperfect competition can occur in labor markets when there is discrimination or when companies or labor unions gain significant market power.

[4] Classic early studies of the human capital model include J. Mincer (1958) and T. W. Schultz (1961). Contributions to hedonic price indexes can be found in Z. Griliches (1971) and S. Rosen (1974).

number of each type of worker reflects the opportunity costs of acquiring a given amount of education or training.[5]

This study, like all previous studies of labor composition, cross-classifies hours by *education, work experience,* and *sex*.[6] There are several surveys which collect data on the educational attainment and earnings of individuals. However, data on training are rarely available. On-the-job training is therefore not a practical characteristic for partitioning workers. Mincer (1974) addressed this problem by developing a wage model which uses work experience rather than training. Hours are also partitioned by sex because many studies indicate that the amount of training per year of work experience may differ by sex (Ben-Porath, 1967; Weiss and Gronau, 1981; Mincer and Polachek, 1974; Sandell and Shapiro, 1980).[7] Appendix D describes completely the collection of the hours data for this study.

To be sure, there are many other earnings differences between workers, and many of these relate to differences in marginal products between workers. Two highly investigated sources of earnings differences are industrial and occupational differentials. In Jorgenson, Gollop, and Fraumeni (1987), a shift of otherwise identical workers between industries paying different wage rates is treated as a change in efficiency due to a reallocation of labor. This source of efficiency change is not included in this study's measures of labor composition and labor input. Instead, this source of efficiency remains in the BLS multifactor productivity measures.[8] Occupational wage differentials have been shown to be significant and persistent even after controlling for education and work experience.[9] Many explanations have been proposed for the cause of these differentials. Some are consistent with a competitive supply of labor, and some are not. Regardless, competitive firms hiring workers will still equate the prevailing wage rate with the value of the worker's marginal product, so occupational differentials should reflect differences in marginal products.

Nevertheless, this study does not partition hours by occupation for several reasons. Occupation and education are highly correlated. Denison argued that education is often a prerequisite for entrance into an occupation and that education increases the ability to work in many occupations.[10] Education therefore has a direct return through higher wages within an occupation and an indirect return through the ability of more educated workers to receive higher wages in other occupations. As a result, disaggregating hours by both education and occupation would limit the earnings differentials by education to the direct returns and understate the total impact of changes in educational attainment on labor composition. Furthermore, the classification of hours by even broad occupational groups is weak in the study's primary data sources.[11] The error rate in assigning workers to occupations is large, so changes in the distribution of workers across occupations or occupational wage differentials include a large random component which would reduce the reliability of the labor composition measures. The errors might contaminate the measure of labor composition if misclassification of workers into occupations is not random.

Measuring the price of labor

Once the hours of work have been appropriately cross-classified, prices for each kind of labor are required. Hourly earnings for each type of worker are based on econometrically estimated hourly earnings functions instead of sample estimates of average hourly earnings. The hours and earnings of each type of worker are taken from the March supplement to the Current Population Survey. For many types of workers, the average hourly earnings are based on just a few observations, and so the variance of these hourly earnings estimates is large. Wage equations can increase the precision of the earnings estimate whenever earnings of different types of workers are structurally related.[12] Furthermore, wage equations provide a method for dividing hourly earnings into payments for each trait used in the partial indexes of labor composition found in appendix H.

While wage models are a novel approach as part of the measurement of labor composition, models of wages have been one of the most intensely investigated relationships in labor economics during the last 20 years. These studies—for example, Mincer (1974), Griliches and Mason (1972), Hanoch (1967), Heckman and Polachek (1974), Mincer and Polachek (1974), Welch (1973), Ashenfelter and Johnson (1972), Lewis (1983), Mincer and Ofek (1982), and Sandell and Shapiro (1980)—have investigated various aspects of the wage model, but all have as their core, the concept that education and training enhance the skills of workers. The wage models used in this study are described in detail in appendix

[5] An exception arises in the case of firm-specific training. In such cases, firms and workers share the cost and reward to training, and marginal products rise faster than wages so that firms can recoup their investment. Barron, Black, and Loewenstein (1989) attempt to measure the difference between wages and marginal products for a group of newly hired, disadvantaged youths. They find that productivity rises at twice the rate of wages over the first 2 years of employment with the firm.

[6] Previous studies of labor composition use age rather than experience directly. This study uses a model to estimate the level of experience for each type of worker. The advantage of this approach is that some of the restrictions imposed on earlier models are removed. The experience model is the subject of appendix C.

[7] The issue of discrimination by sex is discussed at length later in this appendix.

[8] Jorgenson, Gollop, and Fraumeni (1987) treat the reallocation of labor across industries as a separate factor in explaining productivity change and so exclude this effect from multifactor productivity.

[9] See Kruger and Summers (1986).

[10] See chapter 4 of Denison, *Accounting for United States Economic Growth 1929-1969*.

[11] The Current Population Survey conducts periodic reinterviews with a subsample of respondents to ascertain the accuracy of the data collection. Reinterviews indicate that more than 10 percent of respondents are reclassified by 1-digit industry and nearly 20 percent are reclassified by 1-digit occupation.

[12] For example, the earnings function developed by Mincer related changes in earnings over a worker's career to changes in training investments. Consequently, differences in earnings by years of work experience must conform to certain restrictions on the parameters of his model.

E. Wage models estimated using the National Longitudinal Survey can be found in appendix B.

The estimated models are fundamentally human capital models of wage determination. In this context, education and on-the-job training (as measured by work experience) develop skills which in turn enhance worker productivity. Because workers are paid the value of their marginal product in long run competitive equilibrium, the coefficients for schooling and experience measure the contribution of these activities to both earnings and worker productivity.[13]

The functional form of the wage model, equation 13, is based on the work of Mincer:[14]

$$(13)\ \ln(W_{ijk}) = a + b * S_i + c * X_j - d * X_j^2 + f * Z_k.$$

The log of the wage, W_{ijk}, is a function of i years of schooling, S, j years of experience, X, and the k^{th} bundle of other traits, Z. First and second order experience terms are included to capture the observed parabolic pattern of earnings with age. This study uses hourly wages to be consistent with labor data measured in hours. Separate equations are estimated for men and women. Appendix E discusses the exact functional form used in this study.

In equation 13, the parameters b, c, and d measure the return to education and experience for each sex holding the effects of other traits constant. While equation 13 could be estimated with all these traits, it is not practical to disaggregate the hours matrix across these other traits because the number of observations per cell would be insufficient for reliable estimates. Consequently, an average wage for each education-experience cell must suffice.

At least two strategies are possible for estimating the wages for each cell of the matrix. One would be to omit the variables represented by Z from the wage equation. In this case, the resulting parameters would represent the return to education and experience for the prevailing distribution of other traits in each year.[15] That is, the return to a high school diploma depends not only on the value of the diploma but also on the proportion of high school graduates that work full time. Since the proportion of full-time workers can change from year to year, the return to a diploma may change not because the wage or return to education for any group of workers actually changed but because the proportion of full-time workers (who on average are paid more per hour than part-time workers) declined. This example shows that workers within an education-experience cell may not be homogeneous. A resultant bias can arise because changes in the distribution of workers erroneously leads to a change in measured labor composition when no such change occurred.

As an alternative strategy, we can estimate the full equation 13, but use only the returns to education and experience and the intercept. The average effect over all other characteristics, $\overline{Z}$, is added to the intercept. Once equation 13 has been estimated, each worker is assigned hourly earnings based on equation 14.

$$(14)\ \mathrm{Ln}(W_{ij}) = a' + b * S_i + c * X_j - d * X_j^2$$

where $a' = (a + f * \overline{Z})$

Since the altered intercept, a', is an average for all persons of the same sex, it can be interpreted as the return to all characteristics other than education and experience. Changes in the altered intercept result from changes in the return to observed traits (f) or changes in the distribution of either observed (Z) or unobserved traits.

Note the effect of equation 14 on the labor composition measures. Shifts in the distribution of hours between men and women will continue to affect the labor composition measures. However, shifts between groups of men (or groups of women) in other traits (Z) no longer influence the measures of labor composition. Since all men are assigned the mean return to other traits, the earnings of all men are affected proportionately. Such a shift leaves compensation shares and ultimately labor composition unaffected. Even though workers within an education-experience cell may still not be homogeneous, changes in the distribution of the Z variables do not affect the compensation shares of workers of different education or experience levels and so can not lead to erroneous changes in measured labor composition. For example, suppose the proportion of new college graduates who work part time increases from one year to the next, but the proportion of new high school graduates who work part time remains unchanged. The mean hourly wage of new college graduates would fall relative to new high school graduates assuming part-time work pays less. Part-time work is one of the variables included in the set of Z variables. Whatever the return to part-time work may be, it is incorporated into the altered intercept. Equation 14 would indicate that the relative earnings of new college to high school graduates remained unchanged and so would not influence the labor composition measures. In effect, equation 14 removes shifts in the part-time/full-time distribution of the work force from the measures of labor composition and labor input. Instead, the impact of the part-time/full-time distribution is included in the measures of multifactor productivity ($\dot{A}/A$).

This approach to labor composition sought to identify a short list of specific characteristics and their effect on productivity change rather than to examine a large number of unidentified labor-related influences. The effects of increases in education, fluctuations in experience, and patterns of female labor force participation are particularly relevant areas of public interest.

If these are the principal sources of labor composition, the resultant labor composition measures will encompass virtual-

[13] See Barron, Black, and Loewenstein (1989) for the exception of specific training.

[14] The form derived by Mincer (1974) uses an approximation of an infinite series of investments to achieve a log linear estimating form. See *Schooling, Experience and Earnings* for a complete derivation.

[15] This is equivalent to the strategy of Jorgenson, Gollop, and Fraumeni of using the average wage rate of all persons within a cell.

ly all of the potential changes in labor composition. Any unidentified changes in labor composition remain part of the MFP measure along with returns to scale, the effects of capacity utilization, changes in work effort, changes in the organization of the work place, etc. *Appendix E reveals that wage differentials between workers are predominantly due to differences in education and experience.* In addition, appendix D shows that shifts in the education distribution and sex composition of workers are the largest changes in the composition of workers over the past 40 years. Substantial labor composition changes are due either to large shifts in the hours of different types of workers or shifts between workers with large wage differentials. Such shifts did take place in the education, work experience, and sex composition of the work force; therefore, one can reasonably conclude that shifts in the other traits did not make an important contribution to labor composition change.

Two additional points should be noted. First, if the parameters on the Z variables are not large, or not strongly correlated with education and experience, measured labor composition is likely to be similar regardless of which strategy is chosen. As will be seen, only the part-time work schedule has both a relatively large coefficient and a high correlation with work experience. So while the choice of methods may have different theoretical implications, from a practical standpoint, partitioning hours by these other traits would likely lead to results similar to those reported in this bulletin.

Second, a more comprehensive measure of labor composition could differ from the results reported in this bulletin if additional traits not included in the wage model strongly influence earnings. As previously explained, a worker's occupation and industry are explicitly excluded from the wage model. Tests presented in appendix B that gauge the effect of occupation and industry on the wage model show that both traits are highly correlated with earnings. Sensitivity of labor composition to the inclusion of occupation and industry are explored in a comparison of labor composition studies found in appendix G.

Alternative interpretations of the wage model

Alternative interpretations of the education and experience coefficients are possible. Education and work experience may provide benefits of value to the firm, but education and work experience may not directly increase a worker's marginal product. Consequently, the link between increased earnings and additional education or work experience may depend on other unobserved causal factors.

The premise that firms have imperfect information concerning an employee's abilities forms the basis for the assertion by screening and signaling models (Spence, 1974, Taubman and Wales, 1973) that education serves as a valuable informational tool to help firms identify the ability or potential for training of workers rather than a direct source of additional skills. Firms are willing to pay more for workers of greater ability but cannot directly observe ability. Instead, education serves as a signal to identify workers of high ability and does not directly increase a worker's ability to perform in a job. Because (firms believe) education is correlated with ability, a firm that chooses a highly educated worker is also more likely to have chosen a worker of greater ability. The apparent relationship between earnings and education is, in this model, a relationship between ability and earnings.

Job matching models have been used to explain the positive relationship between earnings and experience (Jovanovic, 1979a and b, Abraham and Farber, 1987). In these models, workers search for the highest paying job. That job is the one with the best match between a worker's skills and the firm's needs. Workers will change jobs when higher pay is offered, and the likelihood of receiving a better offer depends on the quality of the match between worker and employer. Over a worker's career, wages rise, and it appears that more experienced workers earn more when, in this model, some or all of the increase results from an increasingly good job match over time.

The multiperiod implicit contracts literature (Lazear, 1979 and 1981, Medoff and Abraham, 1980) suggests that earnings rise with experience because both employers and employees have agreed to defer compensation until later in the employee's career. Workers agree to this arrangement because they earn a rate of return on their deferred compensation. Employers use deferred compensation to maximize the lifetime work effort of employees. Experienced workers who provide insufficient effort risk being fired and losing their deferred compensation. Consequently, experienced workers know that their (current plus deferred) compensation with their current employer is greater than what they could earn with a different employer, and so experienced workers have an incentive to provide an acceptable level of work effort. Under such a contract, workers earn less than the value of their marginal product early in their careers and more than their marginal product later.

Even within the human capital model, experience parameters may be understated in the case of employer paid specific training. If firms pay for training, they can recoup their investment by paying workers less than their marginal products. In this case, marginal products rise faster than earnings after training, and the experience parameters would be understated. The importance of this effect depends on the proportion of training which firms pay for and the rate of worker turnover. If workers effectively discard firm-specific training when leaving a firm, the fraction of total training which is firm-financed, specific training can be expected to be small for the average worker.[16]

The screening, job matching, and implicit contract models have a common element: It may be wrong to infer that acquiring additional education or experience will raise a worker's marginal product. If any of these models is correct, a different or more detailed classification of hours would be required be-

[16] Barron, Black, and Lowenstein (1989) attempt to measure changes in worker productivity separately from changes in earnings. For a sample of newly hired workers, they find that productivity measures rise twice as fast as earnings over the first 2 years of employment.

cause labor is not homogeneous within each education/experience group. The implicit contracts model goes one step further and argues that observed earnings differentials overstate differences in marginal products. Specific training paid for by firms also leads to a discrepancy between earnings and marginal products, but in this case earnings understate marginal products. In either case, some adjustment to the estimated parameters would be needed.

The measures of labor composition can differ using these alternative models. Continuing to use education and experience rather than the true underlying factors (as suggested by the screening and search models) would lead to different estimates of labor composition change unless the distribution of hours classified by these underlying traits changes at the same rate as the distribution of hours by education and experience. Alternatively, if labor composition is still intended to measure changes in the distribution of education and experience, the screening, job matching, and implicit contracts models—but not models of specific training—imply that the measured earnings differences between education groups or between experience groups are overestimated and should be reduced from those obtained in the wage model of equation 11.[17] Although the exact magnitude of the adjustments to earnings required by these models is unknown, sensitivity tests found in appendix G measure the impact on the labor composition measures of altering wage differentials by education and experience.

Market imperfections. Workers are assumed to be paid the value of the marginal product of their labor. However, compensation may not strictly reflect the marginal product of labor. In this case, compensation shares become inappropriate weights in the Tornqvist index of labor input because compensation no longer measures the flow of labor services from an hour of work. The Tornqvist index, in fact, incorporates the input's marginal product into the relevant elasticity for use as its weight for each type of labor. Compensation shares can substitute for a direct measure of elasticities because compensation is assumed to reflect a worker's marginal product. When labor market imperfections are suspected, a distinction between the value of a worker's marginal product and compensation should be made, and the former would ideally be used in measuring labor composition.

As indicated earlier, there are more sources of wage differentials than can be measured with reasonable precision. Since the direct measurement of marginal products is impractical, the analysis must be limited to only those market imperfections which may cause the greatest distortion in the labor composition measures.

Labor market discrimination is a widely noted market imperfection. Discrimination against blacks and ethnic minorities is assumed to take the form of employee discrimination.[18] This form of discrimination largely redistributes earnings between sets of workers rather than between workers and firms. Although earnings for any group of workers affected by discrimination may not indicate their marginal product, earnings of all workers should closely approximate the marginal product of all workers. As a result, the education and experience coefficients for neither whites nor minorities reflect their productivity, but the coefficients for the combined sample reflect the average productivity of all groups.[19] In this study, a single common set of education and experience coefficients are estimated from a pooled sample of workers of all races.[20] This is consistent with the assumption that wage differentials by race are the result of discrimination.

There is less agreement in the literature concerning discrimination against women. The inability to measure skills directly confounds analysis of this form of discrimination. Most studies of labor composition (Jorgenson, Gollop, and Fraumeni (1987); Denison (1985); and Chinloy (1981)) are forced to use some variant of a worker's age to measure work experience or investments in training. An advantage of this study is the construction of measures of work experience for men and women. As a result, earnings differences between men and women can be decomposed into differences in the quantity of experience and education and the returns to these and other traits. In this study, the quantity and return to a worker's experience and education are regarded as measures of the worker's underlying skills. As a result of using experience rather than age in the wage models, the estimated male/female earnings differences explained by factors other than education and work experience are smaller than in many other studies.

The smaller coefficients (especially for work experience) for women can be interpreted as either labor market discrimination or less acquired skills. Even if men and women invest in skills at the same rate, the average workweek for employed women is nearly 6 hours less than for men, and women thus

[17] The net effect of these criticisms may be small because they have opposite effects on the measures of labor composition. These models suggest that the earnings differentials by both education and experience are overstated. Given the increasing educational attainment of the work force, a smaller earnings differential would lead to a smaller measure of labor composition. Declining levels of work experience combined with a smaller earnings differential by experience would lead to a higher estimate of labor composition growth.

[18] See Becker (1971) for a discussion of employee discrimination and its impact on wages.

There are, of course, other models of discrimination. These include employer discrimination (Becker, 1971), statistical discrimination (Arrow, 1971), occupational segregation (Bergmann, 1973), and dual labor markets (Cain, 1975). Each of these models leads to a specific method for measuring the difference between earnings and marginal products.

[19] Firms continue to hire workers until marginal costs equal the marginal product of labor. Under employee discrimination, however, firms may pay white workers a premium in excess of their marginal product. The wage paid to minorities is their marginal product less the marginal cost to firms of additional premiums paid to white workers. However, the sum of all premiums and penalties is zero, and by averaging over all workers, firms pay all workers their marginal product.

[20] This is also the approach used by Jorgenson, Gollop, and Fraumeni (1987), and Denison (1985), in their studies of labor composition.

have fewer hours at work to acquire skills through on-the-job training.[21] It is widely acknowledged that expected labor market absences for women result in a smaller optimal investment in skills acquired by on-the-job training (Weiss and Gronau, 1981, Sandell and Shapiro, 1980). Furthermore, labor market absence (Mincer and Ofek, 1982) may result in a faster erosion of skills. Expected labor market absences have also been used to construct models of occupational self-selection (Polachek, 1974). All these arguments imply that smaller work experience coefficients reflect less accumulation of skills rather than labor market discrimination.

These arguments suggest other causes of earnings differentials, but they do not preclude labor market discrimination. In addition to the forms of discrimination cited above, the comparable worth literature (Treimann and Hartmann, 1981) argues that the work of women may have been systematically undervalued. Most empirical studies of male/female earnings differences (Beller, 1982; Corcoran and Duncan, 1979; Malkiel and Malkiel, 1973; and Oaxaca, 1973) after controlling (imperfectly) for labor market absences and other male/female differences still find that worker characteristics cannot fully account for this earnings differential. However, the proportion of the wage gap accounted for by observable differences in traits varies from 0 percent (Blinder, 1973) to more than 70 percent (Malkiel and Malkiel, 1973).

Because of the uncertain extent of discrimination in labor markets, the sensitivity analysis of appendix G uses alternative assumptions consistent with varying degrees of sex discrimination. These alternatives imply that women have more skills than their earnings suggest, and the growth rate of labor composition would be higher. Appendix G demonstrates that unless all earnings differences result from discrimination, the measures of labor composition would be trivially higher in the absence of discrimination.

Marginal products may also diverge from earnings if unions and firms excercise market power. In addition, the payment of compensating differentials for a variety of reasons including risk of injury, length of workweek, or risk of unemployment may create earnings differentials between workers who differ only in their work schedule or type of firm.[22]

The impact of some of these possible market imperfections as well as the issue of measuring the returns to ability separately from education are considered further in appendix B. However, in part for practical reasons, hours are not partitioned by any of these characteristics. Based on the estimated coefficients of the effects found in appendix B, the impact of excluding these earnings differences from the labor composition model is believed to be negligible.

[21] In 1986, female employees who were at work in the March survey week and worked at least 1 week last year worked 35.6 hours per week. The comparable workweek for men was 41.3 hours.

The data also indicate that women work nearly 2 weeks less per year than men. Women work 44.8 weeks, and men work 46.6 weeks per year.

[22] See Brown (1980) for a survey of empirical studies of compensating differentials.

This study uses wages excluding benefits. Workers are not indifferent to how compensation is divided between wages and benefits. Worker preferences over the two forms of compensation are further complicated because wages are taxable but benefits generally are not. Consequently, the education and experience coefficients may differ between a model of wages and a model of compensation. However, Kiker and Rhine (1987) develop a wage model which includes some aspects of both working conditions and benefits. They find that the education and experience coefficients are only slightly larger when nonwage benefits are included.

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Appendix B. Experiments with the National Longitudinal Survey

As indicated in chapter II, the prices of labor used to determine the weights in the Tornqvist labor input index are based on wage equations estimated annually from the March Supplement to the Current Population Survey (CPS) and from the decennial censuses of 1950 and 1960. The CPS was used because it is an ongoing annual survey; it is selected on a representative basis; the information collected has remained fairly consistent over time; and it has been available for much of the period studied. There are, however, some potentially serious problems in using the CPS to estimate wage equations. First, it does not collect information on actual work experience, so researchers using this body of data have had to rely on proxies for experience. The proxies generally used are age or "potential" work experience (age minus 6 minus years of schooling) which measures the number of years after leaving school. Second, the CPS does not collect information on several relevant variables. Omitting these variables could significantly bias the estimated contribution of education and experience to the earnings measures.

This appendix presents some tests of these two issues using the National Longitudinal Survey of Labor Market Experience (NLS). The NLS data used in these tests are divided into four cohorts, each consisting of approximately 5,000 individuals.[1] The younger cohorts comprise men ages 14 to 24 in 1966 and women ages 14 to 24 in 1968. The older cohorts comprise men ages 45 to 59 in 1966 and women ages 30 to 44 in 1967. Each cohort has been followed continuously since the beginning of the survey, and the participants have been questioned on their schooling, work history, and personal background both during and prior to the survey period.[2] Thus, the NLS collects information on actual work experience as well as data for computing potential work experience. It also includes statistics on a number of variables (IQ scores, current job tenure, etc.) that the CPS does not collect.

The remainder of this appendix is composed of three sections. The first section compares the effects of using potential rather than actual work experience in the wage equations and discusses the implications for the estimated prices of labor. The second section analyzes the effects of omitting selected variables—that is, relevant variables reported in the NLS but not in the CPS—on the estimated wage equations. The final section summarizes the results of the tests that are described in the appendix.

Potential versus actual work experience

Conceptually, human capital wage models need to measure the flow of productive services from cumulated investment in on-the-job training, but this information is not presently available and may never be available. The Mincer human capital earnings function discussed in appendix A and used in this study attempts to capture these service flows by assuming that the rate of investment in on-the-job training is linearly related to actual work experience.[3] Thus, using potential work experience as a proxy for actual work experience requires the additional assumption that the two are "close substitutes."

Econometrically, the use of potential work experience in the wage function introduces an errors-in-variables problem.[4] This tends to bias the coefficient of the linear term for experience toward zero. The coefficients for schooling and other variables may also be biased if they are correlated with the measurement error in experience, but it is difficult to know magnitude or direction. In any case, the magnitude of the bias is determined by how close a substitute potential experience is for actual experience.[5]

Researchers working with human capital wage models have generally found that potential work experience is an appropriate proxy for actual experience for men, but not for most women (that is, women who have had interrupted work experiences because of child bearing and rearing). One of the more convincing studies of this issue is by Malkiel and Malkiel (1973) who had access to a microdata source that enabled them to compute alternative measures of experience. They found that for both men and women the estimated coefficients with respect to the first degree experience terms were lower when potential rather than actual experience was used. For

[1] In 1979, a new cohort was added to the NLS project. It consisted of about 13,000 men and women who were between the ages of 14 and 21 at the start of 1979.

[2] For a description of the NLS data, see *The National Longitudinal Surveys Handbook* (1982). Although the NLS is another possible source of data for estimating wage equations, it was not used because (1) it is obviously not a representative sample of the U.S. labor force in any given year, (2) it may suffer from attrition bias because discontinued respondents may not be representative of the entire sample, and (3) the NLS fails to solicit the same information annually, so that there are a number of "breaks" in the variables over time.

[3] Mincer (1974), pp. 83-89.

[4] An early attempt to refine the measures of potential experience can be found in Hanoch (1967) who measured actual average age of labor force entry for each level of education.

[5] Let $P_i = E_i + u_i$ where P_i is potential experience, E_i is actual experience, and u_i is a random error of measurement. The magnitude of the bias in the experience coefficient is determined by the ratio of the variance of u_i to the variance of E_i. If P_i is a close substitute for E_i, then the ratio will be close to zero and the bias will be small. For a lucid discussion of the errors-in-variables problem, see Rao and Miller (1971), pp. 179-84.

men the difference is small, whereas for women it is substantial. They also found that the estimated schooling coefficients were larger for both males and females when potential rather than actual experience was used, but the numerical values of these biases were about the same for both sexes. Garvey and Reimers (1980) using NLS data also found that the returns to schooling for women were higher when potential experience rather than actual experience is used in earnings functions.

Measurement of NLS Work Experience. At the time the tests in this appendix were conducted, the NLS data provided up to 14 years of continuous work history as well as information on work experience for years prior to the initial survey. The measure of "actual" work experience derived from these data and used in the analysis below is, in fact, a hybrid of potential and actual experience.[6] In addition, because of inconsistencies in the data, the estimates were constructed using information that implied the greatest amount of work experience. Thus, to the extent that the measures of actual work experience are biased, they are biased in the direction of reducing the difference between actual and potential experience.

For those years prior to the initial survey year, work experience was initially measured as the time elapsed between the year of the first reported job (or, if not available, year of birth plus years of completed schooling plus 6) and 1 year prior to the initial survey. From this, any reported time not working just before the first survey and absences from the labor force for child bearing and rearing were subtracted. No experience prior to the first survey was assigned to persons who specifically stated that they had "never worked before" the birth of a child.

During the survey period, respondents reported the number of weeks worked, weeks unemployed, and weeks out of the labor force "in the past year." The NLS did not always survey the respondents at exactly the same date each year, and so time in and out of the labor force did not always sum to 52 weeks. It was assumed that the correct response was the one implying the most work experience. The NLS does not survey respondents in all years. In some years, questions covering work experience in a nonsurvey year were asked at a later date. The responses indicated which interval (for example, 0-13 weeks) appropriately described the work pattern. In such cases, the midpoint of the interval was used. When no information was available, it was assumed that the omitted work history was identical to that of the preceding year.

Fortunately, some of these estimates could be checked by reports of total weeks worked in a given 5-year period. When these figures were available, they were used if the estimate implied more work experience than the sum of the 5 individual years. In addition, job histories through 1976 or 1977 are available, but the incomplete reporting of this data limits its usefulness. Nevertheless, work experience as determined by job history was used whenever it implied greater experience than the two preceding estimates.

Comparisons of regression results. Hourly earnings equations were estimated for each of the four cohorts; one set was based on potential experience and the other on the measure of actual experience described above. In order to emphasize the contribution of experience to earnings, survey data for 1980 were used for younger men; 1976 data for older men; 1981 data for older women; and 1980 data for younger women. Three kinds of comparisons were carried out; they involved the estimated coefficients, coefficients of determination, and the estimated peak earnings years.

Table B-1 compares the estimated constant terms, schooling and experience coefficients in the earnings functions, for the four cohorts using potential and actual work experience. (The complete specification of the earnings function and parameter estimates can be found in table B-7 of this appendix.) As indicated earlier, the use of potential work experience introduces an errors-in-variables problem in the experience variable.[7] This tends to bias all of the estimates in the earnings function, the coefficient of the linear term for experience toward zero, and the others in unknown directions.

The biases are likely to be most pronounced for older women because potential experience is probably the poorest proxy for actual experience for this group. When potential experience is used for this cohort, the constant term is higher, and each of the schooling coefficients is lower than when actual experience is used. In addition, the coefficients for the first and second degree experience terms have the wrong signs and are statistically insignificant. These results occur because women in this age group (44-58) will tend to be on the relatively flat portion of the age-earnings profile. When actual work experience is used, both experience coefficients have the correct signs and are statistically significant. The results are essentially the same for the younger women cohorts; the main difference is that the estimated coefficient for the second degree experience term is not statistically significant when actual experience is used.

As expected, potential experience is a reasonably good proxy for actual experience in the case of younger men. When potential experience is used for the younger male cohort, the estimated constant term and schooling coefficients are close to those based on actual experience. The two experience coefficients have the correct signs, are statistically significant, and, in fact, are numerically about the same as those based on actual experience.

The results for the older male cohort, those ages 55-69, are mixed. The estimated constant term which reflects the earnings of the base group, those with 0-8 years of schooling, is negative and statistically insignificant when potential experi-

[6] For a similar attempt to estimate work experience, particularly work-experience intervals, from the NLS data, see Mincer and Polachek (1974).

[7] There is also an errors-in-variables problem in using the NLS measure of work experience. This is because the NLS measure is, as described earlier, a hybrid of actual and potential experience. The presumption is that the problem is more significant in the case of potential experience than in the case of the NLS measure of work experience, especially with respect to older women.

Table B-1. Estimated coefficients for schooling and work experience in hourly earnings functions using potential and actual experience[1]

Coefficient	Women Age 26-36		Men Age 30-40		Women Age 44-58		Men Age 55-69	
	Potential	Actual	Potential	Actual	Potential	Actual	Potential	Actual
Constant	1.068	0.602	0.851	0.785	1.534	0.695	-0.937	0.257
	(6.7)	(4.7)	(9.0)	(9.0)	(3.9)	(7.9)	(1.1)	(.9)
Schooling								
9-11 years	.145	.203	.173	.178	.158	.164	.066	.133
	(1.8)	(2.3)	(3.0)	(3.4)	(3.0)	(3.5)	(1.5)	(3.2)
12 years	.320	.338	.300	.292	.343	.357	.188	.266
	(3.5)	(4.1)	(5.3)	(6.0)	(6.7)	(8.4)	(4.4)	(7.0)
13-15 years	.406	.418	.420	.393	.466	.511	.330	.420
	(4.2)	(4.8)	(7.0)	(7.8)	(8.3)	(11.0)	(5.6)	(7.8)
16 years	.597	.707	.587	.534	.551	.619	.406	.515
	(5.9)	(7.7)	(9.5)	(10.0)	(8.6)	(11.0)	(5.3)	(7.3)
17 or more	.651	.790	.630	.558	.782	.822	.565	.686
	(6.0)	(8.1)	(10.0)	(11.0)	(12.0)	(14.0)	(7.4)	(11.0)
Work experience								
Experience	.0027	.0521	.0653	.0811	-.0122	.0346	.1079	.0410
	(.1)	(2.4)	(5.3)	(7.6)	(.5)	(5.7)	(2.7)	(2.8)
Experience squared	.00006	-.000004	-.00163	-.00212	.00016	-.00047	-.00140	-.00054
	(.1)	(.0)	(3.9)	(5.6)	(.5)	(3.2)	(3.2)	(2.8)

Source: National Longitudinal Survey: 1980 for Women age 26-36, 1980 for men age 30-40, 1981 for women age 44-58, and 1976 for men age 55-69.

[1] The estimated earnings functions also include race, Hispanic origin, veteran status, marital status, number of dependents, residence status, health problems, region, and part-time work schedule.

ence is used. It is also statistically insignificant when actual experience is used. This reflects the small number and wide dispersion of earnings of individuals within the group. When potential experience is used, the estimated schooling coefficients are lower than when actual experience is used. The returns to an additional year of potential experience are greater for older men than for younger men. The relative returns are reversed when actual experience is used. Since men's earnings usually peak before age 55, the high return to potential experience for older men seems unlikely. Thus, potential experience also appears to be a dubious proxy for the older male cohort.

The second comparison examines the coefficients of determination found in table B-2. For the female cohorts, the coefficients of determination are much higher when actual rather than potential experience is used. For both of the male cohorts, the goodness of fit is virtually the same whether the function is estimated with potential experience or with actual experience. These findings are consistent with the expectation that potential experience is a suitable proxy for actual experience for men but not for women.[8]

The third comparison involves the estimated year of peak earnings based on potential and actual work experience. The earnings function includes a second degree polynomial for experience that is supposed to capture the commonly observed parabolic earnings-experience profile that rises to a peak and then declines. The estimates of peak earnings years based on the earnings functions for the cohorts in the NLS are presented in table B-3.

The estimated peak earnings years for the two male cohorts based on potential work experience are close to those using actual experience. In both of the cases, the differences amount to 1 year or less. In contrast, the estimated earning functions for women based on potential experience do not even yield peak earnings years; for the younger female cohort no peak year is implied, and for the older female cohort, the equation implies a minimum rather than a peak earnings year. The result for the older female cohort is more tenable when actual experience is used. These results provide further evidence that potential experience is not an appropriate proxy of actual experience for women. Potential experience would seem to be an especially inappropriate proxy for older women, who are more likely to have had interrupted work histories.

In sum, the comparisons presented in this section are generally consistent with those reported by other researchers working with earnings functions: The use of potential work experience as a proxy for actual work experience results in biased estimates, and the biases are especially large for women and

[8] The inclusion of job tenure, which is not collected by the CPS, and unionization, which has generally not been collected by the CPS, in the earnings function raises the value of all of the coefficients of determination. Including job tenure also results in the fits of the equations with potential experience being very close to those of the equations with actual experience, even for women. The values of the $\bar{R}^2$'s are as follows:

	Potential experience	Actual experience
Younger women	0.22	0.24
Younger men	.28	.29
Older women	.33	.35
Older men	.33	.32

Table B-2. Coefficients of determination ($\bar{R}^2$) for hourly earnings functions using potential and actual work experience[1]

Cohort	Potential	Actual
Women ages 26-36 in 1980	0.14	0.20
Men ages 30-40 in 1980	.25	.26
Women ages 44-58 in 1981	.25	.31
Men ages 55-69 in 1976	.28	.27

[1] To obtain $\bar{R}^2$, the coefficient of determination R^2 is adjusted for degrees of freedom. Values of $\bar{R}^2$ are presented instead of values of R^2 so comparisons can be made to the results given in footnote 8 in the body of this appendix.

Table B-3. Estimated year of peak hourly earnings based on potential and actual work experience

Cohort	Years of experience	
	Potential	Actual
Women (age 26-36)	([1])	([1])
Men (age 30-40)	20.1	19.1
Women (age 44-58)	([2])	36.9
Men (age 55-69)	38.5	38.1

[1] No peak year implied by the estimated earnings function.
[2] Minimum rather than peak year implied by the estimated earnings functions.

older men but not for younger men. Thus, potential work experience may be an adequate proxy in earnings functions for younger men only. Although experiments were not conducted with age as a proxy for actual work experience, comparisons reported by Blinder (1976) indicate that it would probably be an even poorer proxy than potential experience.

Effects of omitted variables

The CPS, which is used to estimate wage equations, omits a number of variables that are correlated with years of schooling and experience, a problem that was noted earlier. These omissions could significantly bias the estimated coefficients in the CPS earnings function and, hence, the estimated wages.[9] The NLS includes some of these variables. The purpose of this section is to use earnings functions estimated from NLS data to test the significance of the biases resulting from these omitted variables. In these tests actual experience as described in the previous section appears in the earnings functions.[10]

An example may help illustrate how omitted variables affect the earnings equation and the interpretation of the labor composition measures. In general, persons of greater ability attend school longer than persons of lesser ability. In addition, when both groups have the same amount of education, more able workers should earn more than less able workers. A wage model which includes both education and ability measures should in theory be able to measure separately the contribution of ability and education to earnings. A wage model which only includes education will measure the contribution of both ability and education to earnings. Because education and ability are correlated, the estimated education parameter will include the return to education plus a portion of the contribution of ability. If ability is correlated with other variables in the model, another portion of the contribution of ability will be included in these estimated parameters. To keep this example simple, suppose ability is only correlated with education. The remainder of the contribution of ability will be included in the intercept.

Consequently, it is possible to interpret the education parameters as the return to education and some portion of ability even when ability is not explicitly included in the wage model. The labor composition measures would then also measure changes in education and ability.

Changes in the distribution and reward to ability are unobserved in the above example. As a result, labor composition may change despite no changes in the observed characteristics. For example, the return to education appears to have risen sharply since the 1970's and so has added to labor composition growth. Should this be interpreted as a higher return to education or ability? If it is a higher return to education, encouraging additional schooling would add substantially to labor input and output growth. If it is a higher return to ability, encouraging additional schooling would contribute much less to labor input and output growth. Consequently, the implications of labor composition changes become more murky when variables are omitted from the wage models.

Of course, no wage model can include a comprehensive set of explanatory variables, but by examining the impact of omitting some of the most important variables, one may gain some understanding of their influence on the estimated returns to education and work experience. This may serve as a guide in interpreting changes in measured labor composition.

Nonwage compensation. Conceptually, the dependent variables in the earnings functions should include all kinds of labor compensation. Unfortunately, the CPS and NLS include only wage and salary incomes and, as far as is known, there is no survey reporting labor compensation or that could be matched with the NLS or CPS data.[11]

In order to glean some information on the possible effects of omitting nonwage compensation, data were obtained on nonwage labor cost by occupation and industry from the BLS Employment Cost Index (ECI) for 1980 and were related to

[9] The problem of specification bias due to the omission of a relevant explanatory variable is discussed in most intermediate and advanced econometrics texts; see, for example, Kmenta (1986).

[10] The NLS data were also used to compare the results employing the Gompertz curve instead of the parabolic form for the earnings experience profiles; the fits were slightly better using the parabolic form, which is the form used in this bulletin.

[11] The Survey of Income and Program Participation was specifically designed to collect comprehensive measures of income for a nationally representative sample. It has only recently become available to researchers outside the Bureau of the Census. Future research is planned to examine the usefulness of this survey.

the NLS data.[12] That is, two labor cost measures were constructed: The first includes nonwage benefits "matched" to NLS participants by occupation, and the second includes nonwage benefits "matched" to NLS participants by major industry.

The estimated earnings equations using these two labor cost measures as dependent variables were virtually identical to those obtained omitting nonwage benefits. The small difference due to nonwage benefits by either occupation or industry produced only a scalar adjustment (that is, a change in the estimated constant term) in the original earnings function. These scalar adjustments cancel out in computing the compensation share weights for the Tornqvist labor input index.

Thus, adjusting for nonwage labor cost on the basis of available data would not significantly change the labor input measures. It would, however, be useful to have surveys (or matching surveys) that reported labor compensation including benefits; wages as a percentage of total labor compensation range from 69 percent to 77 percent across industry and occupation, and these variations might well correlate significantly with years of schooling and/or experience. Additional discussion in appendix E, however, indicates that benefits, when available, do not seem to alter significantly the returns to education or experience.

Alternative functional forms of the wage model. Within the human capital model, several alternatives to the basic model are worth noting. First, the productive value of education may be exaggerated because persons of greater innate ability can derive a higher return to schooling. There will necessarily be a positive correlation between schooling and ability, and this implies that omitting ability from the wage model yields education and experience coefficients which are too large.[13]

Empirical evidence indicates that innate ability correlates positively with years of schooling so that higher earnings paid to persons with more schooling reflect, to some degree, payments for ability.[14] Depending on how ability is viewed within the wage model, a resultant bias may also be reflected in the prices of labor which are estimated from the earnings functions using CPS data.

The NLS provides information on IQ scores for the two younger cohorts, but not the older two.[15] In addition, there were no IQ scores reported for about half of the sample of the younger cohorts which may also result in a selectivity bias.[16] One of the major problems in using IQ scores to measure ability is that they may also measure educational background and cultural integration. In an attempt to minimize this problem in the NLS data, the earliest available scores (those prior to 1969) were used in the experiments.

When IQ scores (or IQ deciles) were included in the earnings functions for younger males and females, all of the estimated schooling coefficients that were statistically significant fell (table B-4). For the younger men, those estimated schooling coefficients were reduced by between 0.016 and 0.052 when IQ scores were included, and for younger women those estimated schooling coefficients were reduced by between 0.051 and 0.100. The estimated coefficients with respect to work experience were negligibly affected when IQ scores were included.

Table B-4. Estimated coefficients for schooling and experience in hourly earnings functions including and excluding IQ scores[1]

Coefficient	Women Age 26-36		Men Age 30-40	
	Including	Excluding	Including	Excluding
Constant	0.221 (1.1)	0.713 (4.8)	0.596 (5.9)	0.818 (12.0)
Schooling				
0-8 years	-.462 (.6)	-.498 (.6)	.159 (1.2)	.166 (1.2)
12 years	.104 (1.6)	.124 (1.9)	.132 (3.7)	.148 (4.2)
13-15 years	.191 (2.6)	.242 (3.4)	.169 (4.4)	.200 (5.3)
16 years	.386 (4.8)	.473 (6.1)	.302 (7.3)	.345 (8.9)
17 or more	.441 (4.9)	.541 (6.3)	.391 (8.8)	.443 (11.0)
Work experience				
Experience	.0750 (2.7)	.0763 (2.7)	.0790 (6.7)	.0775 (6.6)
Experience squared	-.00093 (.6)	-.00093 (.6)	-.00222 (4.2)	-.00216 (4.1)
Omitted variable				
IQ Score	.0051 (3.6)	-	.0023 (3.1)	-

Source: National Longitudinal Survey: 1980 for women age 26-36 and 1980 for men age 30-40.

[1] Earning functions excluding IQ score are not estimated for the identical sample as the earnings functions in other tables. Here, the sample is limited to those reporting IQ scores. Older male and female samples did not report IQ scores.

The results for younger males are roughly the same as those reported by Griliches (1977) which were based on more refined data sets and econometric techniques. The somewhat larger differences for younger women may reflect the quality

[12] For a description of the BLS Employment Cost Index, see chapter 8 of the *BLS Handbook of Methods*, Bulletin 2414.

[13] The effect of omitted variables on the remaining coefficients depends on both the direction and strength of the correlation between the omitted and included variables. See J. Johnston (1973), pp. 168-69.

The measurement of ability is very difficult. Ability is multidimensional, and a variety of measures have been used in different studies. Its inclusion in wage models has led to a wide range of estimates of the relative importance of ability and education. See Griliches and Mason (1972), Hause (1972), and Taubman (1976) which suggest the education coefficients are overstated by 10-45 percent.

[14] See, for example, Griliches and Mason (1972) and Taubman (1976).

[15] A number of questions have been raised about the use of IQ scores as measures of ability. However, these are the only ability measures available in the NLS.

[16] The reduced sample used in the regression analysis includes virtually no one with less than 9 years of schooling.

of the NLS data. Because of the uncertainty with regard to the extent of the biases and the possibility that they may be small, it was decided that *ad hoc* adjustments for differences in ability would not be made in calculations of the prices of labor.

The omission of ability from the wage model is not as serious as it might seem. As indicated above, ability may exert an independent effect on earnings. On the other hand, ability may have no direct effect but may instead increase the return to education. That is, more able persons can derive more from schooling than less able persons. If so, the education coefficient can be re-interpreted as the productive value of both the direct effect of schooling and the indirect effect of ability on the effectiveness of schooling. The labor composition index would then measure the change in both effects.

To further expand the interpretation of the parameters, the impact of a year of schooling or experience on earnings and productivity can differ across individuals. Schooling quality can vary across time as well as between individuals (Johnson and Stafford, 1973; Welch, 1966 and 1973; Bishop, 1989). The incentive and pattern of investments in on-the-job training also varies across individuals. (Ben-Porath, 1967; Weiss and Gronau, 1981; Sandell and Shapiro, 1980)

The quality of individual schooling varies because of differences in the student's motivation or the school's resources. The schooling coefficient should then be viewed as the average return to all persons within the given schooling group or the return for the average quality of schooling. Similarly, the experience parameters measure the return to acquired training averaged over persons of diverse levels of training. The coefficients for years of schooling and work experience in this study should therefore be viewed as an average of all persons at each level.[17]

Job Tenure. In the human capital model, the rising portion of the earnings-experience profile is explained by investment in on-the-job training that the employee pays for at least in part; this can be either general training which the employee pays for entirely or firm-specific training for which the employer and the employee share the costs.[18] Theoretically, the layoff rate should be lower for employees with specific training because the layoff would involve a capital loss by the employer; the quit rate should also be lower since quitting would mean a capital loss for the employee.[19] Because of these considerations, economists have associated job tenure with specific training.[20] This means that we might expect the relationship of earnings to job tenure to differ significantly from the relationship of earnings to total experience which includes both specific and general on-the-job training.

The NLS reports information on current job tenure for each of the four cohorts. The earnings functions were estimated with and without job tenure, entered linearly (table B-5). In each of the four cohorts, the estimated coefficient for job tenure was statistically significant, and the magnitude of the coefficient was considerable. Except for younger women, the coefficient with respect to the first degree experience terms was reduced somewhat in each cohort when job tenure was included in the earnings function. In addition, all of the schooling coefficients were somewhat lower when job tenure was included. These results suggest that the estimated prices of labor might be improved if job tenure could be included in the earnings function.

Unionization. There is extensive literature on the possible effects of unionization on labor productivity. Unionization could affect the returns from schooling and/or the shape of the earnings-experience profile, cause firms to seek higher quality workers, or provide a collective voice to discuss working conditions and production methods.[21] Marginal products may also diverge from earnings if unions and firms exercise market power.[22] In most years unionization has not been reported in the March supplement to the CPS. The NLS reported both union membership and coverage by collective bargaining agreements for the participants in the four cohorts. For each cohort, earnings functions were estimated with and without a dummy variable for union membership.

There is no generally consistent pattern of changes in the estimated schooling coefficients when union membership is included in the earnings function (table B-6). In addition, the changes are small except for the older male cohort for which including union membership substantially increases the schooling coefficients for those groups with 13 years or more of schooling. The estimated coefficients with respect to both the linear and second degree experience terms are only negligibly changed when union membership is included; this is true for each of the four cohorts. The dummy variable for unionization is statistically significant in each cohort.[23] It is difficult at this point to judge the net effect of including union-

[17] Changes in schooling quality over time are captured within the model by the use of annually estimated prices for education. If the average level of schooling quality is rising over time, the coefficients will increase to reflect the increasing quantity of skills acquired by the same number of years of schooling. The same effect for work experience is also captured in moving parameters.

[18] These distinctions are due to Gary Becker (1975) who defines firm-specific training "as training that has no effect on the productivity of trainees that would be useful in other firms," (p. 26) whereas "general training is useful in many firms besides those providing it...." (p.19).

[19] Empirical evidence also shows a negative relationship between quit rates and specific training; see Parsons (1977).

[20] Job tenure is also related to unionization so that including it in the earnings function captures some of the effects of unionization on earnings. Tests related to unionization are reported in the next section.

[21] See Ashenfelter and Johnson (1972) for a discussion of unions and worker quality; H. Gregg Lewis (1983) for a survey of union effects; and Stafford and Duncan (1980) for an analysis of the interaction of unions and working conditions. Recent models of unions suggest that unions may have a positive rather than negative impact on productivity, and therefore union wage differentials result from higher marginal products of unionized workers. See Addison and Barnett (1982) for a survey of this issue.

[22] See Heywood (1986) as an example of a study estimating the impact of firm concentration on earnings.

[23] Experiments were also done with union concentration ratios and sales concentration ratios by industry, but the results were either statistically insignificant or untenable.

Table B-5. Estimated coefficients for schooling and experience in hourly earnings functions including and excluding job tenure[1]

Coefficient	Women Age 26-36		Men Age 30-40		Women Age 44-58		Men Age 55-69	
	Including	Excluding	Including	Excluding	Including	Excluding	Including	Excluding
Constant	0.570 (4.5)	0.602 (4.7)	0.848 (9.7)	0.785 (9.0)	0.662 (7.8)	0.695 (7.9)	0.333 (1.1)	0.257 (.9)
Schooling								
9-11 years	.194 (2.2)	.203 (2.3)	.165 (3.2)	.178 (3.4)	.165 (3.8)	.164 (3.5)	.118 (2.9)	.133 (3.2)
12 years	.312 (3.8)	.338 (4.1)	.270 (5.6)	.292 (6.0)	.349 (8.5)	.357 (8.4)	.237 (6.4)	.268 (7.0)
13-15 years	.401 (4.6)	.418 (4.8)	.379 (7.6)	.393 (7.8)	.500 (11.0)	.511 (11.0)	.394 (7.5)	.420 (7.8)
16 years	.678 (7.5)	.707 (7.7)	.514 (10.0)	.534 (10.0)	.595 (11.0)	.619 (11.0)	.481 (7.0)	.515 (7.3)
17 or more	.773 (8.1)	.790 (8.1)	.534 (10.0)	.558 (11.0)	.791 (14.0)	.822 (14.0)	.657 (10.0)	.686 (11.0)
Work experience								
Experience	.0559 (2.7)	.0521 (2.4)	.0700 (6.5)	.0811 (7.6)	.0319 (5.4)	.0346 (5.7)	.0346 (2.4)	.0410 (2.8)
Experience squared	-.00128 (1.1)	-.000004 (.0)	-.00194 (5.2)	-.00212 (5.6)	-.00055 (3.8)	-.00047 (3.2)	-.00049 (2.6)	-.00054 (2.8)
Omitted variable								
Job Tenure	.0349 (6.8)	-	.0120 (6.9)	-	.0138 (9.3)	-	.0103 (8.6)	-

Source: National Longitudinal Survey: 1980 for Women age 26-36, 1980 for men age 30-40, 1981 for women age 44-58, and 1976 for men age 55-69.

[1] The estimated earnings functions also include race, Hispanic origin, veteran status, marital status, number of dependents, residence status, health problems, region, and part-time work schedule.

Table B-6. Estimated coefficients for schooling and experience in hourly earnings functions including and excluding union membership[1]

Coefficient	Women Age 26-36		Men Age 30-40		Women Age 44-58		Men Age 55-69	
	Including	Excluding	Including	Excluding	Including	Excluding	Including	Excluding
Constant	0.588 (4.4)	0.602 (4.7)	0.785 (9.1)	0.785 (9.0)	0.707 (8.1)	0.695 (7.9)	0.189 (.6)	0.257 (.9)
Schooling								
9-11 years	.216 (2.5)	.203 (2.3)	.162 (3.1)	.178 (3.4)	.159 (3.4)	.164 (3.5)	.126 (3.1)	.133 (3.2)
12 years	.352 (4.3)	.338 (4.1)	.277 (5.8)	.292 (6.0)	.357 (8.5)	.357 (8.4)	.282 (7.5)	.268 (7.0)
13-15 years	.449 (5.2)	.418 (4.8)	.399 (8.0)	.393 (7.8)	.510 (11.0)	.511 (11.0)	.466 (8.8)	.420 (7.8)
16 years	.732 (8.0)	.707 (7.7)	.556 (11.0)	.534 (10.0)	.606 (11.0)	.619 (11.0)	.593 (8.4)	.515 (7.3)
17 or more	.792 (8.2)	.790 (8.1)	.575 (11.0)	.558 (11.0)	.785 (14.0)	.822 (14.0)	.747 (12.0)	.686 (11.0)
Work experience								
Experience	.0534 (2.5)	.0521 (2.4)	.0759 (7.2)	.0811 (7.6)	.0331 (5.5)	.0346 (5.7)	.0381 (2.6)	.0410 (2.8)
Experience squared	-.00017 (.1)	-.000004 (.0)	-.00191 (5.1)	-.00212 (5.6)	-.00046 (3.1)	-.00047 (3.2)	-.00049 (2.6)	-.00054 (2.8)
Omitted variable								
Union Membership	.216 (5.5)	-	.129 (6.8)	-	.148 (4.4)	-	.220 (7.2)	-

Source: National Longitudinal Survey: 1980 for Women age 26-36, 1980 for men age 30-40, 1981 for women age 44-58, and 1976 for men age 55-69.

[1] The estimated earnings functions also include race, Hispanic origin, veteran status, marital status, number of dependents, residence status, health problems, region, and part-time work schedule.

ization on the estimated prices, but its effect on the returns to education and experience appear small for most workers.

Summary and conclusions

The prices of labor used to calculate the weights in the Tornqvist labor input index are based on earnings functions using CPS data. (The estimation of these prices is discussed in appendix E.) There are a number of problems in using the CPS data for this purpose. Two of the more important are: (1) the Survey does not collect data on actual experience, so researchers using the series have had to employ a proxy based on potential experience (age minus 6 minus years of schooling) or age; and (2) the CPS omits a number of variables that are correlated with years of schooling and/or experience, and omitting these variables could significantly bias the calculated prices. This appendix reports a number of experiments directed to these questions using the NLS which collects some of the relevant information.

Regarding the use of potential experience rather than actual experience, the results indicate that potential experience, on the whole, is an inappropriate proxy for actual experience. The equations for the female cohorts in which actual experience appears fit the data much better than those in which potential experience appears. For the female cohorts all of the experience coefficients are statistically insignificant when potential experience is used, whereas three of the four coefficients are significant when actual experience is used. For all of the groups except the younger male cohort, the schooling coefficients exhibit a downward bias when the equations are estimated with potential experience. In sum, the use of potential experience as a proxy for actual experience is likely to significantly bias the prices of labor and hence the weights in the Tornqvist labor input index. Because of this, estimation equations for actual experience are developed in appendix C, and the resulting measures are used in the estimation of earnings functions in appendix E.

Turning to the omitted variables, including IQ scores in the earnings functions of the younger cohorts causes nearly all of the schooling coefficients to decrease. Because the available information is limited (in particular there is no ability measure available for the older cohorts), it is difficult to assess how excluding ability from the earnings function affects the prices of labor. Inclusion of an ability measure is likely to lower the returns to education and ultimately the growth rate of the Tornqvist index of labor input (assuming the distribution of ability remains constant over time). On the issue of job tenure, inclusion in the earnings function lowers the coefficient on the first-degree experience term somewhat, with one exception. Additionally, almost all of the schooling coefficients decline, though the differences are generally small. While the omission of job tenure would appear to have the greatest impact on the wage model of all the additional variables considered in this appendix, its effect on the measures of labor composition may be small. The total return to experience (job tenure plus actual experience) may not differ greatly from the return to work experience when job tenure is omitted. Finally, when unionization is included in the wage equation, nearly all of the schooling coefficients change, with the coefficients for the higher education levels tending to rise. However, except for those groups with 13 or more years of schooling within the older male cohort, the differences are again small. Overall, the omission of unionization is not likely to have a substantial effect on the weights used to calculate the Tornqvist labor input index.

Table B-7. Basic estimated hourly earnings functions by cohort, National Longitudinal Survey data[1]

Variable	Women (Age 26-36)	Men (Age 30-40)	Women[2] (Age 44-58)	Men (Age 55-69)
Constant	0.602 (4.7)	0.785 (9.0)	0.695 (7.4)	0.257 (0.9)
Schooling				
9-11 years	.203 (2.3)	.178 (3.4)	.164 (3.5)	.133 (3.2)
12 years	.338 (4.1)	.292 (6.0)	.357 (8.4)	.268 (7.0)
13-15 years	.418 (4.8)	.393 (7.8)	.511 (11)	.420 (7.8)
16 years	.707 (7.7)	.534 (10)	.619 (11)	.515 (7.3)
17 or more	.790 (8.1)	.558 (11)	.822 (14)	.686 (11)
Work experience				
Experience	.0521 (2.4)	.0811 (7.6)	.0346 (5.7)	.0410 (2.8)
Experience squared	-.000004 (.0)	-.00212 (5.6)	-.00047 (3.2)	-.00054 (2.8)
Other characteristics				
Black	-.018 (.4)	-.169 (5.9)	-.082 (2.5)	-.163 (3.2)
Hispanic	.121 (1.2)	.018 (.5)	.035 (.5)	-.087 (.5)
Veteran	-	.056 (3.1)	-	.058 (1.9)
Married, spouse present	.092 (2.1)	.160 (5.6)	.051 (1.1)	.454 (4.7)
Married, spouse absent	.153 (3.2)	.120 (3.5)	.091 (1.9)	.279 (2.6)
Number of dependents[3]	-.057 (4.7)	.027 (3.8)	.005 (.7)	-
Resident of central city	.177 (4.7)	.183 (8.1)	.163 (7.8)	.195 (5.3)
Resident, balance of city	.114 (3.3)	.219 (11)	-	.167 (5.0)
Health problem	.030 (.5)	-.195 (5.9)	-.085 (2.9)	-.144 (4.1)
Southern region	-.097 (3.2)	-.085 (4.7)	-.040 (2.9)	-.113 (3.6)
Part-time worker	.051 (1.6)	-.017 (.3)	-.184 (8.1)	-.250 (5.1)
Degrees of freedom	1463	2460	1578	1338
R^2	.20	.26	.31	.27

Source: National Longitudinal Survey: 1980 for women age 26-36, 1980 for men age 30-40, 1981 for women age 44-58, and 1976 for men age 55-69.

[1] Note that the earnings function used in this appendix contains variables not included in the earnings function described in appendix E (for example, race, number of dependents).

[2] Women age 44-58 use one dummy variable for all residents of an SMSA.

[3] For all women, the number of children is substituted for the number of dependents.

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Appendix C. The Estimation Equation for Work Experience

The Current Population Survey (CPS), the basic data source used to estimate the annual prices for the labor input index, does not collect information on work experience. Potential experience (age minus 6 minus years of schooling) and age, which are available in the CPS, have been used as proxies for actual experience, but they can yield biased estimates of prices of labor.

Appendix B briefly reviewed econometric issues and also presented comparisons of regressions performed with measures of potential and actual work experience, using National Longitudinal Survey of Labor Market Experience (NLS) data. The comparisons suggest that the contributions of education and experience to earnings are both biased when potential experience is used. Biased estimates were found for each of the four cohorts reported in the NLS (that is, for young men, young women, older men, and older women), but they were most pronounced for women. This reflects the fact that the work histories of women tend to show interruptions because of child bearing and rearing. This appendix describes the development of an experience measure which overcomes some of the shortcomings of using potential experience or age in an earnings function.

The prices in the labor input index are obtained in two stages. In the first stage, experience equations were estimated separately for men and women based on actual work histories through 1973. These equations relate potential experience and other information available in the CPS to actual work experience. In the second stage, estimated actual experience and data reported in the CPS on schooling and other worker characteristics are used to estimate earnings functions for men and women. These earnings functions are then used to calculate the prices employed in the Tornqvist labor input index.

The experience equations are the first stage in a two-stage process of estimating earnings functions. This appendix describes the estimation of the experience equations. Appendix E describes the earnings functions. The experience equations are estimated from a matched sample of the Current Population Survey, Social Security Administration (SSA) records, and Internal Revenue Service data.[1] Continuous quarterly work histories (hereafter, referred to as SSA experience or actual experience) from 1937 through 1972 reported to SSA were matched to a variety of worker characteristics available in the March 1973 CPS. The 1973 exact match file is the only one now available for use outside of the Census Bureau.[2]

The body of this appendix is divided into three parts. The first section describes the match sample data base. The second section discusses the specification of the experience equations and presents the estimates. In the final section, the results are tested against alternative experience measures in estimating earnings functions and prices of labor for 1973.

Exact match file

The 1973 exact match file produced by the Bureau of the Census and the Social Security Administration linked the Current Population Survey with Social Security and Internal Revenue Service records. The starting point was the March 1973 CPS. Attempts were made to match these individuals with any or all of their SSA earnings, SSA benefits, and 1972 IRS records. More than 90,000 individuals were matched with their SSA records. These records provide quarterly statistics of continuous work histories covered by the Social Security program for the period 1937-72.

Further adjustments were made to the exact match sample in order to make it more consistent with the scope of the BLS multifactor productivity measures. First, the sample was limited to paid employees 14 years of age or older who were currently employed in private nonagricultural industries, satisfied the SSA-CPS exact match rules, and had only one Social Security number. Workers excluded from the Social Security program until the early 1950's (for example, employees in welfare and religious organizations, domestic workers, agricultural and rail transportation workers) were excluded from the sample because of lack of information on their work experience prior to that time. Black and Hispanic workers were not included in the sample, because they were more likely to have been employed in industries not covered by Social Security.

The SSA-CPS exact match sample omitted information that was available in the CPS and needed for estimating the experience equations. For example, it excluded the number of children who were ever born to women, which was available in the CPS. The exact match sample was rematched with the

[1] U.S. Department of Health, Education and Welfare, Social Security Administration, "Studies from Interagency Data Linkages."

[2] There is a 1980 exact match file, but it is not currently available for use outside of the Census Bureau and does not contain as complete a set of variables as the 1973 file. In addition, the Survey of Income and Program Participation collects measures of work experience for a nationally representative sample of the labor force, and it has been matched to the Social Security Administration's continuous work history file. BLS is currently examining these data sources, and if any proves useful in providing an updated experience equation, it will be incorporated in future revisions of the BLS labor composition measures.

Table C-1. Summary statistics for exact match file, 1973

Variable	Men	Women
	Mean	
SSA work experience (quarters)	69.52	45.56
CPS potential experience (years)	20.24	19.78
CPS potential experience prior to 1937 (quarters)	4.56	4.34
Age (years)	38.21	37.59
Education (years)	11.86	11.69
	Percent	
Education		
0-4 years (S0-4)	1.09	0.58
5-8 years (S5-8)	11.47	8.06
9-11 years (S9-11)	18.92	18.25
12 years (S12)	39.85	52.79
13-15 years (S13-15)	15.28	14.42
16 years (S16)	8.93	4.62
17 or more years (S17+)	4.45	1.29
Nonprofit industries	5.96	20.90
Ever married	-	76.70
Spouse present	-	58.84
Spouse absent	-	17.86
Never married	-	23.30
Children ever born to women[1]		
No children (KID0)	-	40.67
1 Child (KID1)	-	15.84
2 Children (KID2)	-	20.04
3 Children (KID3)	-	12.26
4 or more Children (KID4UP)	-	11.20
Veteran	49.81	-
Sample size	17,136	8,685

[1] Figures based on exact match sample for women with 25 or fewer years of potential work experience and the mean number of children ever born from the 1970 decennial census for all other women.

NOTE: The percent distributions may not sum to 100 due to rounding.

original CPS March 1973 data in order to include these omitted characteristics. The final sample used comprises about 8,700 women and about 17,100 men. Summary statistics and acronyms for common variables used throughout this appendix can be found in table C-1.

There are several additional qualifications in using the exact match sample that should be noted. First, despite attempts to exclude all workers not covered by Social Security, some remain in the sample. Workers in nonprofit organizations were frequently covered by private pension plans and were exempted from Social Security coverage. Such exempt workers were mainly in service industries that are largely nonprofit (such as the medical and educational industries). The exempt workers could not be identified and excluded from the sample. The only way to remove most of those exempt workers from the sample was to exclude those industries from the analysis. Since this would have removed many covered workers from the sample as well, this was not done. An interaction term between potential experience and employment in industries with a large percentage of nonprofit establishments was included in the experience equation to account for the exemption from Social Security coverage of some workers in nonprofit industries.

In addition, it was not possible to determine which of the currently covered workers were previously employed in uncovered sectors because only the current occupation and industry of the worker are reported. Thus, work experience in previous uncovered employment could not be credited to these workers. Table C-1 reveals that men spent more than 85 percent of their potential working lives employed in covered industries. The remaining 15 percent includes periods of unemployment and time out of the labor force, leaving relatively little time for employment in uncovered industries.

Also, in the matched sample, experience is truncated for those persons who worked prior to the inception of the Social Security program in 1937 so that their experience is understated. To take account of this, a variable for "previous experience" was created that equals the number of quarters between the person's leaving school and the first quarter of 1937; persons who left school after 1936 are assumed to have had no earlier experience. That is, previous experience equals the number of quarters of potential experience that the person accumulated prior to 1937. Finally, the reported SSA quarterly figures may not reflect different durations of employment during the quarter; the worker employed only 1 week during the quarter and the worker employed during the entire quarter may be both credited by the SSA for one quarter's employment.

Experience equations

The equations used to estimate actual SSA work experience are essentially empirical in that no attempt was made to develop a theoretical model.[3] Since the purpose of the equation is to estimate actual experience from data reported annually in the CPS, the explanatory variables are limited to those that are consistently reported in the CPS.

Actual quarters of work experience reported to the SSA are modeled as a function of potential experience (age minus years of schooling minus 6), a set of schooling dummy variables, the interaction of potential experience and schooling variables, potential experience prior to 1937, the interaction of potential experience and a dummy variable for employment in a service industry with a high proportion of nonprofit workers, and other worker characteristics. The experience model is linear in form and appends an independent and identically distributed error term. Tables C-2 and C-3 contain a complete specification of the experience model. Each of these variables has been discussed above except for data on marital status and fertility of women. These are discussed below.

[3] Work experience in the current period depends on prior work experience among other factors. A model of work experience which reflects this effect could be modeled by a hazard function (a series of equations in which the probability of current employment depends on those variables found in table C-1 and prior experience). Total experience at a point in time would then be the sum of a set of equations from the point of labor force entry until the present. Such a model of total experience would require data not available in the matched data set.

Note also that the work experience equation covers work histories through 1973. Secular changes in labor force participation especially for women may have substantially altered the parameters of the work experience equation in recent years. If so, measures of work experience after 1973 may be inaccurate and bias the labor composition measures. Tests in appendix G indicate that labor composition measures do not suffer from the use of a single experience equation.

Estimates. Table C-2 shows three estimated experience equations for quarterly SSA cumulated work experience for men. In equation 1, the dependent variable comprises all reported SSA experience; it includes all reported work experience including work experience while attending school. Equation 2 includes the same variables as equation 1 except that the dependent variable includes only work experience after leaving school. To obtain figures for work experience after leaving school, it was assumed that persons attended school continuously and that age upon leaving school is the number of school years completed plus 6. Equation 3, like 2, includes only work experience after leaving school as the dependent variable but excludes the dummy variables for years of schooling.

The coefficients of determination (R^2) are between 0.84 and 0.87, indicating that each equation fits the data fairly closely, particularly when compared with those generally observed in similar household cross-section studies. The coefficients of determination are slightly larger for equations 2 and 3 than for 1. This reflects the fact that the dependent variable (SSA postschool experience) in equations 2 and 3 refers to experience after leaving school, whereas in equation 1 the dependent variable comprises SSA total experience.

Interpretation of the estimated coefficients should be made with caution because little structure was imposed on the experience model. Inclusion of the schooling dummy variables and the interaction of these dummy variables with potential experience are included to permit both the intercept and the rate of acquiring work experience to differ across different schooling groups. Differences between schooling dummy variables may reflect differences in work experience while still in school (especially in equation 1), but also reflect differences in the likelihood of unemployment and spells out of the labor force.

All of the estimated schooling coefficients in equations 1 and 2 are statistically significant (at the 5-percent level of significance). In equation 2 the schooling coefficients for men with 12 or more years of schooling are all positive and lie between 2.00 and 3.30. Each of these coefficients represents the additional quarters of SSA postschool experience that men with at least a high school diploma have acquired in excess of the experience of men with 9-11 years of schooling (when both groups have zero potential experience).[4] For men with 9-11 years of schooling, estimated SSA postschool experience when potential experience is zero is given by the constant, which equals 0.31 and which is not statistically significant. These estimates suggest that men with at least 12 years of schooling tended to acquire 2 or 3 quarters more of SSA experience during their working lives than men with 9-11 years of schooling (all other things equal). The coefficients for the two lowest schooling levels in equation 2 are negative.

The effect of a high school diploma on experience is more pronounced in equation 1 than in equation 2. The difference

Table C-2. Estimated experience equations for men, 1973

Variable	Equation 1	Equation 2	Equation 3
Constant	3.00 (7.1)	0.31 (.8)	2.12 (6.8)
Potential experience	3.85 (84)	3.88 (89)	3.90 (90)
Potential experience squared	-.021 (17)	-.019 (16)	-.021 (18)
0-4 Years schooling	-19.23 (5.2)	-16.78 (4.8)	- -
5-8 Years schooling	-13.15 (12)	-10.66 (10)	- -
12 Years schooling	4.99 (14)	3.21 (9.5)	-
13-15 Years schooling	9.90 (16)	2.01 (3.4)	-
16 Years schooling	14.38 (18)	2.26 (2.9)	-
17+ Years schooling	15.89 (14)	3.17 (3.0)	-
Potential experience * S0to4	-.06 (0.6)	-.14 (1.5)	-
Potential experience * S5to8	.15 (4.6)	.07 (2.2)	-
Potential experience * S13to15	-.05 (1.7)	.11 (4.4)	-
Potential experience * S16	-.15 (4.0)	.12 (3.2)	-
Potential experience * S17+	-.45 (8.2)	-.18 (3.5)	-
Previous experience	-.37 (15)	-.43 (18)	-.47 (19)
Potential experience * nonprofit industries	-.33 (15)	-.32 (15)	-.31 (14)
Degrees of freedom	17,120	17,120	17,131
R^2	.84	.87	.86

NOTE: The dependent variable in equation 1 is all reported SSA experience. Equations 2 and 3 use reported SSA experience after leaving school. Persons are assumed to attend school continuously. Age upon leaving school is defined as the number of years of completed schooling plus 6.

In each equation the dependent variable is measured in quarters, as is previous experience. Potential experience is measured in years.

Source: Current Population Survey, Administrative Record, exact match file.

[4] SSA postschool experience can be positive when potential experience is zero because of the manner in which potential experience is calculated. If a man born in 1950 completed college in 1972 and then worked for two quarters in covered employment, he would have zero years of potential experience (age minus 6 minus years of schooling) and two quarters of SSA postschool experience at the end of 1972. This effect may also reflect discontinuities in attending college.

between the schooling parameters of equations 1 and 2 also increases as the level of schooling increases. It is tempting to interpret this result as additional work experience while in school. That is, high school graduates work 4.99 more quarters while in school than men who complete 9-11 years of schooling. While this may explain much of the differences in these coefficients, as cautioned above, other factors such as unemployment and labor market absences may apply as well.

Equations 1 and 2 also include interaction terms between years of schooling and potential experience. The coefficients on the interaction terms for 13-15 years and 16 years of schooling are positive and significant in equation 2, indicating that men with 1-4 years of college acquired more SSA experience per year than those with 1-3 years of high school (that is, the group with 9-11 years of schooling). This may reflect, in part, the fact that less educated workers tend to experience more or longer spells of unemployment than more educated workers. The negative coefficient on the interaction term for 17 or more years of schooling may have occurred because such workers may have been more likely to work in nonprofit organizations (such as private schools and hospitals) and to be covered by a private pension plan rather than Social Security. In equation 1 the coefficient on that term is even more negative; furthermore, the coefficients on the interaction terms between potential experience and 13-15 and 16 years of schooling are negative in the first equation. One possible explanation for this is that better educated older men in the sample may have had less experience in covered employment prior to leaving school than better educated younger men. The younger men may have actually worked more, or they may have just spent more time in jobs covered by Social Security.

The estimated coefficient with respect to previous experience—that is, potential experience prior to 1937—is negative and statistically significant in all three equations. The coefficient is negative because work experience prior to 1937 is not reflected in the SSA measures of experience.

As expected, the interaction term between potential experience and nonprofit industries has a negative coefficient (and is statistically significant) in each of the three equations. Persons in this sector were not in covered employment and instead were under private pension plans. Such people had fewer quarters of SSA experience than would be expected based on their education and potential experience. As a result, the coefficient on the interaction term between potential experience and nonprofit industries is negative.

Table C-3 shows similar experience equations for women. Again, the dependent variable in equation 1 comprises SSA total experience, including jobs held while in school, whereas those in equations 2 and 3 include only postschool experience. In addition to the explanatory variables in the equations for men, the equations for women also include dummy variables for "ever married" and "number of children." Inclusion of these additional variables can reflect the impact of a woman's responsibilities for children on her participation in labor markets.

Table C-3. Estimated experience equations for women, 1973

Variable	Equation 1	Equation 2	Equation 3
Constant	2.20 (2.4)	0.28 (.3)	1.90 (3.1)
Potential experience	4.26 (35)	4.38 (37)	4.25 (37)
Potential experience squared	-.035 (13)	-.037 (13)	-.036 (13)
Ever married	2.00 (2.0)	2.29 (2.3)	2.93 (3.1)
Potential experience * ever married	-1.02 (6.8)	-1.11 (7.5)	-1.13 (7.8)
Potential experience squared * ever married	.017 (5.4)	.019 (6.2)	.019 (6.2)
Potential experience * KID1	-.30 (3.8)	-.29 (3.8)	-.23 (3.0)
Potential experience * KID2-3	-.83 (12)	-.86 (13)	-.82 (12)
Potential experience * KID4UP	-1.14 (13)	-1.18 (14)	-1.18 (14)
0-4 years schooling	-18.33 (2.3)	-16.73 (2.2)	-
5-8 years schooling	-11.13 (4.3)	-9.66 (3.8)	-
12 years schooling	3.98 (3.8)	1.95 (1.9)	-
13-15 years schooling	8.84 (7.2)	3.28 (2.7)	-
16 years schooling	11.47 (6.6)	3.41 (2.0)	-
17 + years schooling	16.39 (4.7)	6.30 (1.8)	-
Potential experience * S0to4	.19 (.9)	.15 (.7)	-
Potential experience * S5to8	.14 (2.0)	.11 (1.5)	-
Potential experience * S12	-.06 (1.4)	-.02 (.5)	-
Potential experience * S13to15	-.45 (8.2)	-.32 (5.9)	-
Potential experience * S16	-.61 (6.6)	-.42 (4.6)	-
Potential experience * S17+	-1.07 (7.2)	-.80 (5.5)	-
Potential experience * nonprofit industries	-.28 (12)	-.28 (12)	-.32 (14)
Degrees of freedom	8,663	8,663	8,675
R^2	.60	.63	.62

NOTE: The dependent variable in equation 1 is all reported SSA experience. Equations 2 and 3 use reported SSA experience after leaving school. Persons are assumed to attend school continuously. Age upon leaving school is defined as the number of years of completed schooling plus 6.

In each equation the dependent variable is measured in quarters, as is previous experience. Potential experience is measured in years.

Source: Current Population Survey, Administrative Record, exact match file.

The "number of children" refers to the number of related children of any age and marital status living in the household or away attending school. For relatively young women, this measure accurately counts the number of children ever born to these women. However, for relatively older women, these children may be adults and no longer living at home. Consequently, the quality of this measure of children erodes as the age of the women in the sample increases. A comparison of these responses with those of the 1970 decennial census (which does not suffer from this problem) indicated that the data are satisfactory for women with 25 or fewer years of potential experience (age minus schooling minus 6). As a result, for women with 26 or more years of potential experience, these data are not used.

Instead, the mean number of children ever born to women with 26 or more years of potential experience from the 1 in 100 sample of the 1970 decennial census is substituted. That is, the proportion of the women with zero, 1, 2 or 3, or 4 or more children ever born substitutes for the corresponding dummy variable for children. In addition, different proportions are used for each of five educational groups of ever married women, and a separate group is used for all never married women.[5]

The experience equations for women do not fit the SSA-CPS data as closely as those for men. In each of the three cases, the equations explain about 60 percent of the variation in actual experience among women compared with more than 80 percent for men.

The pattern of the schooling coefficients for women in equations 1 and 2 is similar to the pattern for men. The coefficients for those with 8 or fewer years of schooling are all negative. The remaining schooling coefficients are positive, with the coefficients in equation 1 exceeding the corresponding ones in 2 due to the inclusion of work experience while in school.

In contrast to the men's equations, the coefficients on the interaction terms between potential experience and years of schooling in the women's equations decline continuously as years of education rise. There are at least two plausible explanations for this. One is that better educated women in the private-sector work force in 1973 may have been more likely than their male counterparts to have previously been in nonprofit industries and not covered by Social Security.[6] Another possibility is that there may have been a correlation between women's education level and husband's income; in the years prior to 1973, this may have resulted in lower labor force participation for more highly educated women.

As with the men's equations, the coefficients on the interaction terms between potential experience and years of schooling for women with 13 or more years of education are lower in equation 1 than in equation 2. The explanation may well be the same as the one suggested for the men's equations: The older women with high levels of education might have spent less time in covered employment prior to leaving school than the younger women with high levels of education did.

The remaining estimated coefficients of the women's experience equations generally have the expected signs. The sole exception involves the "ever married" variable, which has a positive sign in each equation; however, the interaction term between potential experience and "ever married" does have a negative coefficient in each equation. Furthermore, the positive coefficient on the term for the interaction of "ever married" and potential experience squared has two effects on the pattern of women's work experience. First, the depressing effect of marital status on work experience reaches its maximum effect about 30 years after leaving school. Second, after that point, the effect of marital status diminishes somewhat. This may indicate that many older married women return to the work force after raising their children. All of the coefficients in table C-3 pertaining to children have negative signs, with the magnitudes of the coefficients increasing as the number of children rises. This confirms the expectation that the duration of a woman's absence from the labor force increases with the number of children she has. Finally, interacting potential experience and nonprofit industries yields a negative coefficient, as it does in the men's equation.

Testing the alternative experience measures

The second stage of the two-stage approach for obtaining prices of labor involves estimating earnings functions using the estimated values from the experience equations (equation 1 in tables C-2 and C-3) as the measure of experience. This section employs the 1973 data to compare these results (and the results of alternative approaches) against those based on actual SSA total experience which is only available for that year.

Schooling and experience coefficients. Tables C-4 and C-5 show the estimated schooling and experience coefficients for men and women, respectively, based on alternative measures of experience. The estimated earnings functions from which these are taken are shown in tables C-9 and C-10. The first column shows the results based on actual SSA total experience including experience acquired while in school. That is, it is based on the "preferred" measure of experience and, therefore, the one against which all of the other results are compared.

The wage equations used in this appendix differ from those actually used in the construction of the labor composition indexes. Additional explanatory variables are used here to minimize the impact of the experience measures on the estimated parameters. If significant parameter differences remain, this more comprehensive specification of the wage model strengthens the inference that the choice of the experience measure is the source of the parameter differences.

The second column shows the estimated coefficients when only SSA postschooling experience is used to measure experience. Comparing these results with those in column 1 indicates that the estimated returns to additional schooling tend to be higher when postschool experience is used, particularly for

[5] The proportions are based on a sample of women who worked in the previous year in the private business sector. The educational groups are for women with 0-8, 9-11, 12, 13-15, and 16 or more years of completed schooling.

[6] According to table C-1, 21 percent of the women in the sample were classified in the industries with a large percentage of nonprofit establishments compared to only 6 percent for men.

Table C-4. Estimated schooling and experience coefficients of hourly earnings functions based on alternative measures of experience, men, 1973

Variable	Actual SSA experience: Total	Actual SSA experience: Postschool	Estimated SSA experience: Total	Estimated SSA experience: Postschool	Potential experience
	(1)	(2)	(3)	(4)	(5)
Intercept	0.539 (9.5)	0.560 (10)	0.465 (8.0)	0.525 (9.2)	0.538 (9.5)
S5to8	.007 (.1)	.006 (.1)	.003 (.0)	.005 (.1)	.019 (.4)
S9to11	.081 (1.5)	.081 (1.5)	.080 (1.5)	.084 (1.6)	.114 (2.1)
S12	.186 (3.6)	.202 (3.9)	.181 (3.5)	.198 (3.8)	.245 (4.7)
S13to15	.280 (5.3)	.325 (6.1)	.269 (5.1)	.320 (6.0)	.364 (6.8)
S16	.468 (8.7)	.524 (9.7)	.454 (8.4)	.515 (9.5)	.560 (10)
S17+	.577 (10)	.632 (11)	.540 (9.7)	.603 (11)	.644 (11)
Experience	.0116 (20)	.0120 (22)	.0151 (20)	.0142 (22)	.0109 (22)
Experience squared	-.000051 (12)	-.000057 (14)	-.000082 (15)	-.000081 (22)	-.000053 (22)
R^2	.32	.33	.32	.32	.32

Source: Table C-9.

those with 12 or more years of schooling. The higher coefficients are in effect, "compensating" for the omitted work experience while in school. As shown in appendix A, the share weights used to construct the labor input index are based on the product of hourly prices and actual hours worked. Therefore, prices calculated from an earnings function using only postschool experience would impart a positive bias to the weights on the hours of workers with higher levels of schooling. Since there has been a secular shift in hours worked from persons with less to those with more education, the labor input index would include an upward bias over time. This is the reason that prices based on actual SSA total experience are the preferred prices. There is also a slight difference in the estimated experience coefficients when experience acquired while in school is omitted.

Column 3 shows the results based on estimated SSA total experience, that is, based on estimated experience using equation 1 in tables C-2 and C-3. This functional form, estimated annually, is similar to the one actually used to calculate the prices for the labor input index. Comparing these results with those shown in column 4, where estimated SSA postschool experience is used, indicates the same possible systematic bias described above if the less comprehensive measure of estimated experience is used.

As discussed earlier, potential experience is a commonly used proxy for actual experience, and thus it can be considered as an alternative to estimated SSA total experience. Column 5 shows the estimated coefficients based on potential experience.[7] The coefficients indicate that the schooling coefficients based on potential experience are larger than those based on the preferred measure, actual SSA total experience shown in column 1. The estimates based on estimated SSA total experience, shown in column 3, generally lie below those in column 1; also, the estimates in column 3 are for the most part closer to those in column 1 than are those in column 5. That is, the earnings functions based on estimated SSA total experience yield better estimates of the schooling coefficients than those based on potential experience; this is so for both men and women. The higher estimated schooling coefficients based on potential experience reflect, among other things, the omission of experience acquired while in school. Following the same reasoning as above, schooling coefficients estimated from earnings functions based on potential experience would cause the weights in the labor input index to be biased, and the biases would be positively related to years of schooling.

The results with respect to the experience coefficients are somewhat mixed. In the case of men, the estimates based on potential experience are very close to those based on the preferred experience measure; in fact, they are closer than those based on estimated SSA total experience. In the case of women, on the other hand, the estimated experience coeffi-

[7] The earnings function based on potential experience can be viewed as a "partially" reduced form where the coefficients with respect to schooling include both the structural relationship between education and work experience and the net relationship between schooling and earnings.

Table C-5. Estimated schooling and experience coefficients of hourly earnings functions based on alternative measures of experience, women, 1973

Variable	Actual SSA experience		Estimated SSA experience		Potential experience
	Total	Postschool	Total	Postschool	
	(1)	(2)	(3)	(4)	(5)
Intercept	0.413 (3.8)	0.425 (4.0)	0.397 (3.6)	0.422 (3.9)	0.424 (3.9)
S5to8	.080 (.7)	.084 (.8)	.089 (.8)	.093 (.9)	.118 (1.1)
S9to11	.129 (1.2)	.133 (1.3)	.159 (1.5)	.163 (1.5)	.195 (1.8)
S12	.213 (2.1)	.225 (2.2)	.244 (2.3)	.256 (2.4)	.300 (2.9)
S13to15	.335 (3.2)	.364 (3.5)	.360 (3.4)	.387 (3.6)	.429 (4.0)
S16	.470 (4.4)	.512 (4.8)	.489 (4.5)	.527 (4.9)	.569 (5.2)
S17+	.635 (5.6)	.671 (5.9)	.642 (5.5)	.676 (5.8)	.715 (6.2)
Experience	.0093 (11)	.0091 (11)	.0098 (6.9)	.0093 (7.4)	.0063 (9.2)
Experience squared	-.000039 (5.2)	-.000040 (5.4)	-.000065 (4.2)	-.000064 (4.6)	-.000034 (6.0)
R^2	.14	.14	.12	.12	.12

Source: Table C-10.

cients based on both potential and estimated SSA experience differ substantially from those based on the preferred experience measure. The coefficients for the first order term indicate that the coefficients using actual experience and estimated experience are very close, and the potential experience coefficient is too large. Conversely, the second order experience coefficients using actual and potential experience are very close, and the estimated experience term is too large. This pattern suggests that the estimated experience equation will underpredict the wages of older women, and the potential experience equation will overpredict the wages of young (and possibly most) women. One possible but surprising reason for the pattern of the experience coefficients in the potential experience equation can be found in table C-10. While both the estimated and actual experience equations indicate that the number of children has a strong effect on the work experience of women, table C-10 indicates that the coefficient for the number of children in the potential experience equation is effectively zero. One interpretation of this finding is that both the return to experience and the earnings penalty of child rearing are higher than estimated by the potential experience equation. It appears in the potential experience equation that the effects of both work experience and children on wages are combined into the coefficients for potential experience. For measuring labor composition, it appears that women need to be distinguished by family size, but the potential experience equation can not make such a meaningful distinction. For all these reasons, the estimated experience equations appear to better estimate the women's experience coefficients.

Finally, it should perhaps be noted that the coefficients of determination (R^2) for the estimated earnings functions for women are comparatively low, between 0.12 and 0.15. These are about half the size of those for men. However, the low R^2 values are only relevant here if they reflect (1) missing variables that are correlated with years of experience or years of schooling, or (2) a misspecification of the earnings functions, particularly with respect to schooling and experience.

Hourly prices of labor. The above comparisons of the earnings functions are somewhat inconclusive: Although the two-stage approach based on estimated SSA total experience yields less biased estimates of the schooling coefficients than potential experience, it does not do so for the experience coefficients, particularly in the case of men. The net effect can be ascertained by comparing the price estimates, which are based on both the schooling and the experience coefficients.

These comparisons are shown in tables C-6, C-7, and C-8, for men, single women with no children, and married women with two or three children, respectively. Each table is composed of three sections; each section corresponds to a different level of schooling. In the first line of each section, estimates of SSA total work experience corresponding to various levels of potential experience are given. These estimates were calculated with experience equation 1 from table C-2 (for men) and table C-3 (for women). The remaining lines in each section contain estimated prices of labor for workers at a particular level of schooling. The prices on the second line are based on the earnings function which uses potential experience. For each level of potential experience in the tables, the estimated price was computed with the coefficients in column 5 of table C-4 (for men) and table C-5 (for women). The third line of each section contains prices that are obtained with the estimates of experience in line 1 and with the earnings function based on estimated total experience. The coefficients of this earnings function are in column 3 of tables C-4 and C-5. The prices in line 4 of each section are also calculated with the experience estimates in line 1; the earnings function derived with actual SSA total work experience (column 1 of tables C-4 and C-5) is used to calculate these prices.[8]

For men, estimated experience appears to be a modestly better measure. Using the prices in line 4 as the standard of comparison, estimated experience yields better estimates of prices in general for men with 5-8 years or 16 years of schooling. That is, the absolute difference between the price based on estimated experience and the price based on actual experience is, in most of the cases, smaller than the difference between those based on potential and actual experience. For high school graduates potential experience yields slightly better results in the majority of cases. Overall, estimated SSA experience performs better than potential experience. Since men have a more or less continuous attachment to the work force, it should not be surprising that potential experience proves to be almost as strong a proxy as estimated work experience.

[8] There are two ways to calculate prices in line 4. The prices could be estimated using the mean level of actual work experience corresponding to the levels of potential experience or the predicted level of work experience. Since there were not enough observations in the Exact Match File, using mean actual experience for all of groups was not practical and instead estimated experience was used. This might make the prices in lines 3 and 4 more similar than they would be if mean actual experience had been used in the calculations for line 4.

Turning to table C-7, neither potential nor estimated experience seems to be a superior measure for the entire sample of never married women with no children. At lower levels of potential experience, estimated experience yields prices closer to those based on actual experience, while at higher levels, potential experience results in closer estimates.[9] Since never married women are predominantly young, wage estimates for low levels of potential experience are relatively more important.

However, for married women with two or three children, estimated experience is clearly a better measure. In virtually all of the cases in table C-8, the price based on estimated experience is nearer to that based on actual experience than is the price yielded by potential experience. In table C-8, the mean absolute difference between the price based on predicted experience and actual experience is less than 4 cents. The corresponding difference between the price based on potential experience and actual experience is almost 13 cents. These results for married women are not surprising, because one of the primary motives for the use of the two-stage approach is to obtain a more appropriate measure of work experience for women with interrupted work histories.

Tables C-6, C-7, and C-8 underline the problems of trying to find a simple, direct relationship between potential experience and hourly prices of labor, particularly for women. As

[9] Estimated experience may not yield better results for single women in part because they represent a small proportion of the sample.

Table C-6. Estimated work experience (in years) and prices of labor based on alternative measures of experience (in dollars), by schooling and years of potential experience, men, 1973

	Years of potential experience						
	0	5	10	15	20	25	30
5-8 years of schooling							
(1) Estimated experience (years)	0.00	2.15	6.56	10.71	14.59	18.20	22.64
Prices of labor:							
(2) Potential experience	1.75	2.13	2.48	2.77	2.97	3.06	3.01
(3) Estimated experience	1.60	1.81	2.24	2.62	2.92	3.10	3.20
(4) Actual experience	1.73	1.90	2.26	2.56	2.86	3.07	3.25
12 years of schooling							
(1) Estimated experience (years)	2.00	6.68	11.10	15.24	19.12	22.74	26.08
Prices of labor:							
(2) Potential experience	2.19	2.66	3.11	3.48	3.73	3.83	3.77
(3) Estimated experience	2.14	2.69	3.17	3.53	3.75	3.82	3.78
(4) Actual expereince	2.26	2.71	3.13	3.46	3.72	3.89	3.98
16 years of schooling							
(1) Estimated experience (years)	4.35	9.03	13.44	17.59	21.47	25.08	27.29
Prices of labor:							
(2) Potential experience	3.00	3.65	4.26	4.76	5.11	5.25	5.17
(3) Estimated experience	3.18	3.89	4.45	4.83	5.01	5.00	4.90
(4) Actual experience	3.30	3.89	4.41	4.81	5.09	5.25	5.29

NOTE: See text for a description of the methods used to obtain estimated experience and the prices of labor.

Table C-7. Estimated work experience (in years) and prices of labor based on alternative measures of experience (in dollars), by schooling and years of potential experience, single women with no children, 1973

	Years of potential experience						
	0	5	10	15	20	25	30
5-8 years of schooling							
(1) Estimated experience (years)	0.00	3.04	7.88	12.27	16.22	19.73	22.79
Prices of labor:							
(2) Potential experience	1.72	1.92	2.10	2.22	2.29	2.30	2.24
(3) Estimated experience	1.63	1.81	2.08	2.25	2.34	2.36	2.32
(4) Actual experience	1.64	1.82	2.11	2.35	2.53	2.67	2.75
12 years of schooling							
(1) Estimated experience (years)	1.54	6.57	11.15	15.30	19.00	22.26	25.08
Prices of labor:							
(2) Potential experience	2.06	2.31	2.51	2.66	2.75	2.76	2.69
(3) Estimated experience	2.01	2.35	2.59	2.72	2.75	2.72	2.65
(4) Actual experience	1.98	2.32	2.61	2.85	3.02	3.13	3.19
16 years of schooling							
(1) Estimated experience (years)	3.42	7.76	11.65	15.11	18.12	20.69	22.82
Prices of labor:							
(2) Potential experience	2.70	3.02	3.29	3.49	3.60	3.61	3.52
(3) Estimated experience	2.74	3.09	3.33	3.46	3.51	3.51	3.46
(4) Actual experience	2.72	3.10	3.42	3.67	3.85	3.98	4.07

NOTE: See text for a description of the methods used to obtain estimated experience and the prices of labor.

Table C-8. Estimated work experience (in years) and prices of labor based on alternative measures of experience (in dollars), by schooling and years of potential experience, married women with two or three children, 1973

	Years of potential experience						
	0	5	10	15	20	25	30
5-8 years of schooling							
(1) Estimated experience (years)	0.00	1.33	4.16	6.75	9.11	11.24	13.12
Prices of labor:							
(2) Potential experience	1.72	1.92	2.10	2.22	2.29	2.30	2.24
(3) Estimated experience	1.63	1.71	1.88	2.02	2.13	2.22	2.28
(4) Actual experience	1.64	1.72	1.89	2.04	2.18	2.29	2.39
12 years of schooling							
(1) Estimated experience (years)	2.04	4.86	7.44	9.78	11.89	13.77	15.41
Prices of labor:							
(2) Potential experience	2.06	2.31	2.51	2.66	2.75	2.76	2.69
(3) Estimated experience	2.05	2.24	2.40	2.52	2.62	2.68	2.72
(4) Actual experience	2.01	2.21	2.38	2.53	2.66	2.77	2.85
16 years of schooling							
(1) Estimated experience (years)	3.92	6.04	7.94	9.59	11.01	12.20	13.15
Prices of labor:							
(2) Potential experience	2.70	3.02	3.29	3.49	3.60	3.61	3.52
(3) Estimated experience	2.78	2.96	3.10	3.21	3.29	3.35	3.40
(4) Actual experience	2.77	2.96	3.12	3.26	3.37	3.46	3.53

NOTE: See text for a description of the methods used to obtain estimated experience and the prices of labor.

the tables indicate, the problem partly reflects the fact that potential experience is not uniquely related to estimated actual experience. For example, in the case of men, 10 years of potential experience is equivalent to approximately 7 years of estimated actual experience if the individual has only 5-8 years of schooling; however, it is the equivalent of about 11 years of estimated experience if he has 12 years of schooling and of about 13 years of experience if he has 16 years of schooling. Consequently, the higher prices associated with increased education for a man with 10 years of potential experience also reflect increased actual work experience.

The problem is even more complex for women. There is a similar positive association between estimated experience and years of schooling, for a given amount of potential experience, but in addition, the relationship between estimated experience and potential experience differs between single women with no children and married women with children. Again, using the case of individuals with 10 years of potential experience, the figures in tables C-7 and C-8 show that a single woman with no children has more estimated experience than a married woman with children. (In fact, estimated experience for single women with 10 years of potential experience is about the same as for men in the two lower education groups shown in the tables.) As a consequence, at a given level of potential experience, the hourly price of labor for a single woman (derived with estimated experience) is generally higher than that for a married woman with children at all three levels of education. The price differentials between the two groups of women reflect differences in actual work experience.

In conclusion, the two-stage approach using estimated SSA total work experience yields, on net, more satisfactory results than potential experience, the alternative measure directly available from the CPS. For both men and women, the estimated coefficients for education in wage models based on estimated SSA experience are less biased than those using potential experience. The results for the contribution of cumulated work experience are mixed; estimated experience coefficients perform as well for women, but worse for men. Comparisons of estimated hourly prices indicate that estimated experience is at least as good a measure of work experience as is potential experience for men and never married women, and clearly better for married women with children. Also, the estimated experience equations themselves suggest that estimated experience is a better estimator of actual experience than is potential experience; the statistical significance of coefficients pertaining to education (in equations 1 and 2 for men and women of tables C-2 and C-3) and children (in the women's equations) supports the claim that estimated experience is the more appropriate proxy. Using age as the measure of experience was not tested, but, based on studies done by others, age would yield even more biased results than potential experience.[10]

[10] See, for example, Alan Blinder (1976), "On Dogmatism in Human Capital Theory."

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Table C-9. Estimated hourly earnings functions for men, exact match file, 1973

Variable	Actual SSA experience		Estimated SSA experience		Potential experience
	Total	Postschool	Total	Postschool	
	(1)	(2)	(3)	(4)	(5)
Intercept	0.539 (9.5)	0.560 (10)	0.465 (8.0)	0.525 (9.2)	0.538 (9.5)
S5to8	.007 (.1)	.006 (.1)	.003 (.0)	.005 (.1)	.019 (.4)
S9to11	.081 (1.5)	.081 (1.5)	.080 (1.5)	.084 (1.6)	.114 (2.1)
S12	.186 (3.6)	.202 (3.9)	.181 (3.5)	.198 (3.8)	.245 (4.7)
S13to15	.280 (5.3)	.325 (6.1)	.269 (5.1)	.320 (6.0)	.364 (6.8)
S16	.468 (8.7)	.524 (9.7)	.454 (8.4)	.515 (9.5)	.560 (10)
S17+	.577 (10)	.632 (11)	.540 (9.7)	.603 (11)	.644 (11)
Experience	.0116 (20)	.0120 (22)	.0151 (20)	.0142 (22)	.0109 (22)
Experience squared	-.000051 (12)	-.000057 (14)	-.000082 (15)	-.000081 (22)	-.000053 (22)
Hispanic	-.080 (3.1)	-.092 (3.5)	-.137 (5.2)	-.141 (5.4)	-.142 (5.4)
Black	-.149 (8.6)	-.157 (9.0)	-.184 (11)	-.183 (10)	-.179 (10)
Part time	-.214 (11)	-.204 (11)	-.180 (9.1)	-.184 (9.4)	-.204 (11)
Veteran	.016 (1.5)	.010 (1.0)	.028 (2.7)	.025 (2.4)	.028 (2.6)
Ever married	.236 (16)	.227 (15)	.224 (15)	.223 (15)	.235 (16)
Central city resident	.115 (9.1)	.117 (9.3)	.113 (8.9)	.114 (9.0)	.115 (9.1)
Balance of city resident	.137 (12)	.140 (12)	.146 (13)	.146 (13)	.146 (13)
New England Division	.000 (.0)	.001 (.1)	.008 (.4)	.007 (.3)	.005 (.2)
Mid-Atlantic Division	-.003 (.2)	-.005 (.3)	-.002 (.1)	-.003 (.2)	-.003 (.2)
East North Central Division	-.017 (1.0)	-.015 (.9)	-.004 (.3)	-.005 (.3)	-.005 (.3)
West North Central Division	-.101 (4.6)	-.099 (4.5)	-.088 (4.0)	-.089 (4.0)	-.090 (4.1)
South Atlantic Division	-.102 (5.7)	-.103 (5.8)	-.102 (5.7)	-.103 (5.8)	-.105 (5.8)
East South Central Division	-.131 (5.6)	-.134 (5.7)	-.129 (5.5)	-.131 (5.5)	-.133 (5.7)
West South Central Division	-.191 (9.6)	-.190 (9.6)	-.183 (9.2)	-.184 (9.3)	-.187 (9.4)
Mountain Division	-.034 (1.3)	-.032 (1.2)	-.026 (1.0)	-.027 (1.1)	-.029 (1.1)
Degrees of freedom	14,622	14,622	14,622	14,622	14,622
R^2	.32	.33	.32	.32	.32

Appendix D. Constructing Annual Hours of Work Matrices

This appendix documents the construction of annual hours of work cross-classified by all the characteristics which are necessary to implement the measure of labor composition change. The hours of work are assumed to represent the flow of services (or quantity) from a particular category of labor. Construction of these data is essential to measuring labor composition and generates a wealth of detailed data.

For this study there are two sets of matrices: One set for private business and one set for private nonfarm business. The private business sector is the largest aggregate sector for which the BLS constructs measures of multifactor productivity. The output of this sector is defined in terms of the National Income and Product Accounts.[1] Private business output is gross domestic product (GDP) less the output associated with general government (10 percent of GNP in 1985), owner-occupied housing (6 percent in 1985), households and nonprofit institutions (3 percent in 1985), statistical discrepancy (1 percent in 1985), and government enterprises (1 percent in 1985). Therefore, private business output makes up approximately 75 to 80 percent of GDP covering those parts of the economy for which there are independent measures of both output and inputs (capital services and hours). The private nonfarm business sector is private business less agricultural production.

A complete set of data for any sector consists of a matrix for each year, 1948 to the present, containing hours cross-classified by education, sex, marital status, age, and number of children ever born. Table D-1 shows the characteristics and dimensions of the hours matrices. A complete matrix contains 8,064 cells.

However, large groups of cells in each matrix are empty by construction. For example, neither marital status nor number of children is a significant predictor of work experience for men; therefore, the matrix for men is collapsed to exclude these traits. For men there are thus only 504 out of 4,032 cells that could possibly be nonzero. So, the effective matrix is 4,536 cells: 504 for men and 4,032 for women. In addition, there are a negligible number of teenagers with 16 years or more of education. Even exceptions to this would most likely not be part of the private business sector. Nor would one expect to find very many, if any, employed teenage women with more than 3 children. Furthermore, the hours of workers aged 70 or more are small, so many cells for these workers are zero. The number of nonzero cells for each annual matrix ranges between 1,793 and 2,454.

Table D-1. Characteristics and dimensions of annual hours matrices

Characteristics	Dimension	Types
Education	7	Years of schooling 0-4 5-8 9-11 12 13-15 16 17 or more
Age	72	14-85 years old
Sex	2	Women Men
Marital status	2	Ever married Never married
Number of children ..	4	None 1 2-3 4 or more
Total cells	8,064	

Annual hours are constructed from the product of employment, average weekly hours, and average annual weeks. However, no single consistent data source exists from which the matrices can be generated for the entire time period. The development of the hours data entails the use of three different procedures for different periods. For the period 1967 to the present, the matrices are constructed directly from observations collected for the March sample of the Current Population Survey (CPS). For 1950 and 1960 the matrices are generated from observations from the 1 in 100 or 1 in 1,000 samples of the decennial census. The remaining years (1948, 1949, 1951-59, 1961-66) are generated using an interpolation procedure which incorporates published data from the CPS and the 1950, 1960, and 1967 matrices already generated.[2]

The remainder of this appendix describes each of the three methods for calculating the matrices. The next section dis-

[1] See U.S. Department of Labor, Bureau of Labor Statistics (1983), *Trends in Multifactor Productivity, 1948-81*, pp. 35-38.

[2] There are two matrices constructed for 1967 as with other transition years. One is constructed directly from the March 1968 CPS, and one is constructed using the RAS multiproportional matrix interpolation procedure to conform with the previous year's matrix.

cusses the measures for 1967 to the present; the following section discusses the 1950 and 1960 measures; and the third section describes the remaining years along with a description of the interpolation procedure. The final section presents summary measures of hours.

Hours matrices for 1967 to the present

Matrices of annual hours for 1967 to the present are constructed directly from observation data of March samples of the CPS. The CPS is a monthly household survey that has been conducted since 1940. Information is collected for 60,000 to 65,000 households concerning the status of employment of all occupants. Employment responses refer to the week which includes the 12th day of the month. The March questionnaire is supplemented to include information from the previous year including the amount of earnings, the number of weeks worked, and the average weekly hours worked. The weeks worked data refer to the prior year. The average weekly hours data refer to the prior year beginning with the March 1976 survey and to the survey year for the period 1967 to 1976. The CPS (March 1968 to the present) reported hours are classified by worker traits and assigned to the hours matrix for the year prior to the survey year.

Before the information of an individual respondent is included in the calculation, the person must meet selected criteria. Persons selected must, in the present survey period, have a job and also must have worked during the previous year. Government employees and household workers are excluded. However, because there was no satisfactory way to identify employees of nonprofit institutions, such persons probably remain in the sample, though it is unlikely that the distribution of their hours by experience and education differs from others who are in the sample.

Another possible problem is that the CPS does not record the number of children ever born to women who are presently living in a group quarters arrangement, that is, cooperative or shared housing which is not necessarily a family relationship with a head of household. Earnings and Social Security Administration work histories (see appendix C) suggest that the labor market behavior of these women resembles that of women with no children. Hence, all working women living in group quarters are classified accordingly.

A more substantial problem arises with these data because the CPS reports only the number of children born to a woman *who are presently residing in her household or who are away attending school*. As indicated in appendixes C and E, these data are satisfactory for estimating experience equations for women with 25 or fewer years of potential experience (age minus schooling minus 6), but are not satisfactory for older women. Consequently, in years relying in part or in whole on CPS data (those after 1960), these older women are not categorized by the number of children ever born. Instead, the employment and hours of these women are classified by their education, age, and marital status only.[3] The 1950 and 1960 decennial census data on children report all children ever born regardless of residence. Therefore in the matrices, all women in this period continue to be classified by number of children.

Changes in definitions and reporting have caused two distinct breaks in the procedures for constructing the hours matrices from the CPS data. For each break point two matrices are constructed: One corresponding to the new method and one corresponding to the old method. Annual changes in labor composition are measured using pairs of matrices that follow the same methods. This procedure links changes in labor composition across potential breaks in the series without permitting shifts in the distribution of hours due to definitional changes to be included in the labor composition measures.

In 1976, the CPS began reporting average weekly hours for the year prior to the survey and average weeks worked last year as discrete observations. Prior to 1976, average weekly hours were available for the survey week only and weeks worked data were recorded as intervals. As a result, from 1967 to 1975, annual hours are constructed as the product of employment (for those employed in both the survey week and the previous year), average weekly hours (in the survey week), and average weeks worked in the previous year). Because weeks worked are reported in intervals and not as a discrete number, an average for each interval was estimated using the March 1976 CPS, the first March CPS to collect weeks worked as an observation.

Beginning with the 1976 CPS, total hours are defined as the product of employment, average weekly hours in the previous year, and weeks worked in the previous year measured as an observational variable. Hence, the hours matrices are available on this new measure of weeks and average weekly hours from 1975 to the present.

Beginning with the March 1980 survey, the CPS no longer records the educational attainment of 14-year-olds. As a result, the hours matrices for the years 1948 through 1978 contain the hours of persons age 14 and older. Dropping 14-year-olds from the 1979 survey, a second 1978 hours matrix is created. Consequently, matrices from 1978 to the present contain the hours of persons age 15 and older.

Hours matrices for 1950 and 1960

The 1950 and 1960 matrices are created from the *Census of Population, 1950: Public Use Microdata Sample*, and the *Public Use Sample of Basic Records from the 1960 Census*, respectively. The 1950 Public Use Sample was recently coded as machine-readable data files by a joint project conducted by the Bureau of the Census and the Center for Demography and Ecology of the University of Wisconsin.

The 1950 Public Use Sample is derived from a 1-percent sample of households. The final sample was 464,130 households. For each household only one individual known as the sample line respondent was selected to answer additional questions concerning earnings and educational attainment.

[3] These older women are assumed to have the mean number of children for women of their age, marital status, and education. See appendix C for details of assigning means to these women.

From these records BLS selected all observations that met the previously stated industry criteria (no government employees, household workers, or other nonprofit workers). Average weekly hours in the survey week were recorded as a discrete number as was the number of weeks worked in the previous year.

The 1 in 1,000 Public Use Sample, from the 1960 census, was selected to estimate the employment and hours in 1960, yielding over 37,000 observations. Average weekly hours are recorded in intervals for the survey week, and annual weeks are recorded in intervals for the previous year.

For each of the years two matrices are calculated: An employment matrix and average annual hours matrix. The employment matrix is defined as representing the census year (1950 and 1960, respectively), and the average hours matrix is defined as representing the previous year (1949 and 1959, respectively). Using these matrices plus a 1967 average annual hours matrix constructed from the CPS as endpoints, average annual hours matrices for all intervening years are calculated as a linear interpolation of the endpoints. For example, the 1950 average annual hours matrix is equal to 90 percent of the 1949 matrix and 10 percent of the 1959 matrix. The product of the employment and average annual hours matrices yields the total hours matrix.

In order to adjust for the addition of Alaska and Hawaii, two matrices are calculated for 1960. The first matrix excludes Alaska and Hawaii and is used to correspond with all prior year measures. The second includes Alaska and Hawaii and is used to correspond with all succeeding years.

Hours matrices for 1948-49, 1951-59, and 1961-66

Except for 1950 and 1960, there is no single data source for the years between 1948 and 1967 which allows the direct generation of matrices of annual hours cross-classified by the demographic groups necessary to estimate the amount of experience and education for each worker. The available data are from the published tables that were generated from the monthly CPS (usually March) during those years. These tables generally cover employment cross-classified by two or three different characteristics such as age, sex, and marital status or some other combination. For no year is there a set of tables which covers all characteristics. Neither is there any year with information on the number of children born to working women.

This dearth of data for this period required introducing an interpolation method to compute the hours matrices. A multiproportional matrix interpolation method (often referred to as RAS) was selected.[4] This procedure takes as its starting points the distribution in the previously constructed matrices (the 1950, 1960, and 1967 matrices), and, assuming proportionality, adjusts the starting matrix to conform to the distribution of known submatrices where the submatrices are matrices which contain a subset of the known characteristics for the year in question. In this case the submatrices are the published tables referred to above which include some but not all of the information needed.

The derivation of each of the remaining hours matrices is a two-step process. The first step is calculating the annual employment matrix using the RAS procedure. The second step is calculating the hours matrix which is the product of the employment matrix and the average annual hours matrix. The average annual hours matrix is simply the linear interpolation across time of the average annual hours matrices from the census and/or CPS years as described in the previous section.

The actual construction of the employment matrix, using RAS, requires an initial matrix besides the submatrices. Again, the initial matrix is a linear interpolation of two completed matrices. Table D-2 shows how the initial matrix for each of the years is calculated. The determining factor for each year's initialization is the presence of an education component in one of the submatrices. Table D-3 shows the submatrices for each year by characteristics. Matrices in years without an education submatrix are not estimated until matrices in years with the education component are completed. The initial matrix in years without education is a linear average of previously determined matrices in years with a known education distribution. For example, matrices for 1952 and 1957 are first estimated, and then the 1951 initialization is calculated as an average of the 1952 estimated matrix and the 1950 census matrix. Similarly, the 1953, 1954, 1955, and 1956 initializations are weighted averages of the 1952 and 1957 estimated matrices.

In general, for each year there are three to four submatrices

Table D-2. Construction of initialization matrices used with RAS procedure, 1948-67

Year	Data used
1948	1950 census
1949	1950 census
1951	1/2 1950 census and 1/2 1952 RAS matrices
1952	8/10 1950 census and 2/10 1960 census
1953	4/5 1952 RAS and 1/5 1957 RAS
1954	3/5 1952 RAS and 2/5 1957 RAS
1955	2/5 1952 RAS and 3/5 1957 RAS
1956	1/5 1952 RAS and 4/5 19957 RAS
1957	3/10 1950 census and 7/10 1960 census
1958	2/3 1957 RAS and 1/3 1960 census
1959	1/3 1957 RAS and 2/3 1960 census
1961	1/2 1960 census and 1/2 1962 RAS
1962	5/7 1960 census and 2/7 1967 CPS
1963	1/2 1962 RAS and 1/2 1964 RAS
1964	3/7 1960 census and 4/7 1967 CPS
1965	2/3 1964 RAS and 1/3 1967 CPS
1966	1/3 1964 RAS and 2/3 1967 CPS
1967	1967 CPS

Census = employment matrix constructed from census data.
CPS = employment matrix constructed from CPS data.
RAS = employment matrix constructed using RAS multiproportional interpolation procedure.

[4] The biproportional RAS interpolation method is discussed in Bacharach (1965). Extension to the multiproprtional case can be found in Bishop, Feinberg, and Holland (1975). See Frank Gollop and Dale Jorgenson (1980) for an application of the multiproportional RAS method to labor input data.

Table D-3. Characteristics of submatrices used with RAS procedure, 1948-67

Characteristics of employment submatrix	Year																	
	48	49	51	52	53	54	55	56	57	58	59	61	62	63	64	65	66	67
Age by sex for all persons 14 and over, 9 age groups	x	x	x	x	x		x	x									x	x
Age by sex, 11 age groups						x												
Age by sex, 14 age groups									x	x	x	x	x	x	x	x		
Education by sex, 6 education groups	x			x					x									
Education by sex, 7 education groups											x							x
Education by race and sex, 7 education groups													x		x	x	x	
Marital status of women (all women 14 and over)	x	x	x	x	x		x	x		x	x							x[1]
Marital status of women, 14-35 years and 35+ years						x						x	x	x	x			
Marital status of women by race, 14-35 years and 35 + years																x	x	
Marital status, by age, women, 6 age groups									x									
Race and sex for all persons 14 years and over						x	x	x	x	x	x	x	x	x	x	x		x

[1]16 years and over.

which describe the distribution of selected characteristics of the complete matrix. The initial matrix is then adjusted proportionally to each submatrix available as a recursive procedure. After the full initial matrix has been adjusted to each submatrix, the process is repeated. This iterative procedure stops when the largest relative change in any cell of the full matrix is no greater than 1 percent.

Tables D-4, D-5, and D-6 display the final estimates of the proportions of hours at work for all years by selected characteristics. These tables represent, as well as possible, the distribution of the work force (employees, unpaid family workers, and proprietors) in the private business sector of the economy. Excluded are all government employees (Federal, State, and local), household workers, and employees of nonprofit organizations which can be clearly identified in the CPS. This last group of workers mostly includes members of religious organizations.

The tables illustrate the distributions generated by either the RAS procedure or the actual observed data from the CPS or the census. Table D-6, in particular, is displayed only to indicate how the distributions of number of children and education have changed over time for women. The changes displayed here are attributable to a host of factors including changing fertility patterns and a changing age distribution.

Table D-4. Percentage distribution of hours at work in private business by education for men, 1948-90

Year	Years of schooling						
	0-4	5-8	9-11	12	13-15	16	17+
1948	8.3	35.6	20.5	23.1	6.5	3.6	2.4
1949	9.3	36.0	19.5	21.4	7.1	4.0	2.7
1950	9.1	35.9	19.5	21.5	7.2	4.1	2.7
1951	8.8	35.1	19.1	22.4	7.4	4.3	2.9
1952	8.5	34.4	18.8	23.3	7.6	4.4	3.0
1953	8.1	33.4	19.0	24.0	7.7	4.7	3.2
1954	7.6	32.4	19.1	24.7	7.7	4.9	3.4
1955	7.1	31.3	19.4	25.7	7.8	5.1	3.6
1956	6.7	30.3	19.6	26.5	7.9	5.3	3.7
1957	6.2	29.3	19.8	27.3	7.9	5.5	3.9
1958	5.9	28.9	20.2	26.8	8.6	5.5	4.0
1959	5.5	28.4	20.7	26.4	9.3	5.5	4.0
1960	5.2	27.8	21.1	26.1	10.1	5.6	4.1
1961	4.8	25.8	20.4	28.2	10.2	6.5	4.2
1962	4.5	23.8	19.6	30.2	10.2	7.3	4.3
1963	4.1	22.9	19.6	31.5	10.2	7.2	4.5
1964	3.7	22.0	19.6	32.8	10.2	7.0	4.7
1965	3.8	21.1	19.3	33.8	10.1	7.5	4.5
1966	3.4	20.4	19.5	34.4	10.2	7.8	4.4
1967	2.9	18.6	18.8	35.3	11.9	7.6	5.1
1968	2.7	17.9	18.7	35.8	12.2	7.7	5.0
1969	2.5	17.0	17.7	36.4	12.8	8.2	5.4
1970	2.4	15.7	16.9	37.2	13.5	8.6	5.7
1971	2.3	14.8	17.2	37.2	13.8	8.8	5.8
1972	2.3	12.9	16.3	38.8	14.6	9.1	6.0
1973	2.1	12.4	15.7	38.6	15.2	9.5	6.5
1974	1.7	10.7	14.9	38.7	16.0	10.9	7.2
1975	1.8	10.8	14.9	38.9	15.9	10.7	7.0
1976	1.7	10.1	15.0	38.5	16.4	10.9	7.3
1977	1.7	9.7	14.5	38.4	17.4	10.9	7.5
1978	1.5	8.9	13.6	39.0	18.1	11.1	7.8
1979	1.4	8.5	13.7	39.0	17.8	11.5	8.1
1980	1.4	7.8	13.1	39.5	17.8	12.1	8.3
1981	1.3	7.3	12.5	39.6	17.6	12.5	9.2
1982	1.3	6.5	11.7	38.6	18.0	13.8	10.2
1983	1.0	6.4	10.9	39.2	18.4	13.9	10.1
1984	1.1	6.2	11.0	39.2	18.7	14.1	9.8
1985	1.1	5.9	10.5	39.3	19.2	14.5	9.5
1986	1.0	5.6	10.7	38.9	19.3	14.5	10.0
1987	1.1	5.1	10.6	39.3	18.8	14.8	10.2
1988	1.1	5.1	10.2	38.8	19.5	14.7	10.6
1989	1.3	4.9	10.0	38.7	19.9	15.1	10.1
1990	1.3	4.5	9.7	39.1	20.2	15.1	10.1

NOTE: Total may not sum to 100.0 due to rounding.

Table D-5. Percentage distribution of hours at work in private business by education for women, 1948-90

Year	Years of schooling						
	0-4	5-8	9-11	12	13-15	16	17+
1948	4.8	29.5	18.8	36.3	6.8	2.5	1.3
1949	4.5	25.4	20.5	35.5	9.2	3.2	1.7
1950	4.5	25.4	20.3	35.7	9.2	3.2	1.6
1951	4.6	25.8	20.1	36.4	8.4	3.2	1.6
1952	4.6	26.2	19.8	37.0	7.6	3.2	1.6
1953	4.3	25.5	19.9	37.9	7.7	3.2	1.6
1954	4.1	24.7	19.8	38.8	7.8	3.3	1.6
1955	3.7	24.1	19.8	39.6	7.8	3.3	1.6
1956	3.4	23.4	19.9	40.4	7.9	3.4	1.6
1957	3.1	22.5	19.8	41.3	8.0	3.5	1.7
1958	3.0	22.3	20.6	40.2	8.8	3.4	1.6
1959	2.9	22.5	21.7	38.7	9.5	3.2	1.5
1960	2.7	21.7	22.2	38.0	10.6	3.3	1.5
1961	2.5	20.5	20.8	40.3	10.5	4.1	1.2
1962	2.3	18.9	19.4	43.0	10.6	4.8	1.0
1963	2.2	18.5	19.4	44.2	10.1	4.4	1.2
1964	2.1	18.2	19.4	45.2	9.7	4.0	1.4
1965	2.0	17.4	19.2	45.9	9.7	4.2	1.6
1966	1.6	16.7	19.1	47.4	10.3	3.5	1.5
1967	1.3	15.1	18.9	48.0	12.0	3.5	1.1
1968	1.4	13.5	18.6	50.2	11.7	3.3	1.3
1969	1.1	12.4	17.7	50.6	12.3	4.1	1.7
1970	1.1	11.7	17.2	50.3	13.3	4.3	2.1
1971	1.2	10.5	16.9	51.4	13.6	4.8	1.5
1972	1.1	9.5	16.0	52.5	14.2	4.8	1.8
1973	1.0	8.9	15.5	50.8	15.9	5.6	2.2
1974	.9	8.3	14.9	50.8	16.2	6.3	2.8
1975	.8	7.7	15.1	50.3	16.4	6.9	2.8
1976	.9	7.3	14.9	50.4	16.8	7.1	2.7
1977	.8	6.9	14.4	50.2	17.6	7.3	2.8
1978	.8	6.0	13.1	50.4	18.5	7.7	3.4
1979	.7	5.7	12.4	50.2	18.7	8.6	3.6
1980	.6	5.2	11.9	50.0	19.8	8.9	3.6
1981	.7	4.8	11.3	49.8	19.7	9.3	4.4
1982	.6	4.4	10.4	48.7	20.6	10.4	4.9
1983	.6	3.9	9.9	48.4	21.0	11.0	5.2
1984	.6	3.7	9.7	47.3	22.0	11.6	5.1
1985	.5	3.4	9.1	47.0	22.6	12.2	5.2
1986	.4	3.4	8.9	46.7	22.7	12.3	5.5
1987	.7	3.1	9.1	45.8	23.1	12.6	5.6
1988	.7	2.8	9.0	45.6	22.8	13.2	5.9
1989	.6	2.8	8.6	45.0	23.3	13.5	6.3
1990	.7	2.7	8.4	44.3	23.5	14.1	6.3

NOTE: Total may not sum to 100.0 due to rounding.

Table D-6. Percentage distribution of hours at work in private business by education and number of children, for women with 25 or fewer years of potential work experience, 1948-90[1]

Year	Number of children: 1 — Years of schooling: 5-8	12	13-15	16	Number of children: 2-3 — Years of schooling: 5-8	12	13-15	16
1948	3.1	5.6	1.1	0.4	3.8	5.9	1.1	0.3
1949	2.6	5.6	1.5	.5	3.2	5.9	1.6	.4
1950	2.7	5.8	1.6	.5	3.4	6.1	1.6	.5
1951	2.6	6.3	1.5	.5	3.4	7.1	1.7	.5
1952	2.5	6.6	1.4	.5	3.4	8.1	1.7	.5
1953	2.4	7.1	1.5	.5	3.4	9.2	1.8	.6
1954	2.2	7.2	1.5	.5	3.2	9.8	2.0	.6
1955	2.0	7.8	1.5	.5	3.1	11.0	2.1	.7
1956	1.8	8.1	1.6	.5	2.9	11.8	2.3	.7
1957	1.6	8.2	1.6	.5	2.7	12.3	2.4	.8
1958	1.5	8.1	1.8	.5	2.6	12.4	2.7	.8
1959	1.5	8.2	2.0	.5	2.7	12.8	3.1	.8
1960	1.3	7.7	2.1	.4	2.6	12.3	3.3	.8
1961	1.1	8.5	2.1	.6	2.3	12.7	3.2	1.0
1962	1.0	9.1	2.1	.7	2.0	13.2	3.1	1.1
1963	1.0	9.4	1.9	.6	1.9	13.0	2.9	1.0
1964	.9	9.7	1.8	.5	1.9	12.8	2.6	.9
1965	.8	9.8	1.7	.6	1.8	12.5	2.5	.9
1966	.8	10.1	1.8	.4	1.7	12.5	2.6	.8
1967	.7	10.1	2.0	.4	1.5	12.1	2.7	.7
1968	.6	10.1	1.9	.4	1.3	12.8	2.3	.7
1969	.5	10.2	2.0	.5	1.1	12.8	2.4	.7
1970	.6	10.2	2.4	.6	1.0	12.8	2.8	.8
1971	.5	10.1	2.4	.7	1.0	13.8	2.8	.9
1972	.3	11.3	2.6	.7	.8	13.0	3.0	.9
1973	.4	10.2	2.7	.6	.8	13.0	3.0	1.0
1974	.4	10.1	2.9	.8	.8	13.2	3.4	1.2
1975	.4	9.9	3.1	.8	.9	12.9	3.9	1.4
1976	.4	10.0	2.7	.9	.7	13.0	3.7	1.3
1977	.4	10.4	3.0	.9	.8	13.0	3.9	1.4
1978	.3	10.3	3.2	1.1	.6	12.9	4.3	1.2
1979	.3	10.5	3.4	1.2	.7	12.8	4.2	1.6
1980	.3	10.8	3.8	1.2	.6	12.7	4.4	1.8
1981	.3	10.8	3.9	1.3	.6	13.0	4.5	1.7
1982	.2	10.6	3.5	1.7	.6	13.0	4.7	2.0
1983	.2	10.7	3.9	1.9	.5	13.2	4.9	2.2
1984	.2	10.9	4.5	1.8	.4	12.2	5.3	2.3
1985	.2	10.3	4.5	2.0	.4	12.5	5.2	2.2
1986	.2	10.3	4.7	2.1	.6	12.4	5.5	2.3
1987	.2	10.5	4.7	2.0	.5	12.2	5.5	2.5
1988	.2	10.5	4.7	2.0	.4	12.4	5.5	2.6
1989	.2	10.4	4.6	2.5	.5	12.1	5.9	2.6
1990	.2	10.0	4.5	2.5	.4	11.7	6.0	2.6

[1] Years of potential work experience are defined as age less years of completed schooling less six. Accordingly, 25 years of potential work experience corresponds to a different age level for each educational group.

For any year 100 percent is hours of all working women with 25 or fewer years of potential work experience.

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Appendix E. The Estimated Earnings Functions

This appendix describes the estimation of the earnings functions used to calculate the price of labor in the labor input index. These functions relate hourly earnings to years of schooling and experience in addition to other worker characteristics. They are estimated separately for men and women. Earnings functions are not estimated separately by race or ethnicity for reasons discussed in appendix A and because the sample sizes for blacks and ethnic groups are too small to generate reliable estimates for these groups.

The 1950 and 1960 decennial censuses and the annual March supplement to the Current Population Survey (CPS) for each year from 1968 to 1991 are used to estimate earnings functions for the year prior to the survey. (Note that the surveys collect data on the previous year's activity, so the year of the data precedes the survey year.) Although the CPS has been carried out since 1948, data on individual observations for years prior to 1968 are not available, thus earnings functions for the earlier years cannot be estimated. Nevertheless, the longevity and the constancy in collecting essentially the same set of information in the CPS make it possible to estimate nearly identical equations annually beginning with the 1968 survey. For years when neither decennial census nor CPS data are available (1948-49, 1951-59, and 1961-67), the hourly price of labor is obtained by linear interpolation and extrapolation of the estimated coefficients of the earnings function between terminal years. The next section of the appendix briefly describes the decennial census and CPS data, and discusses major issues in using these data to estimate earnings functions. The third section describes the specification of the earnings functions and presents the empirical results. Tests of alternative equations are also discussed in this section. The final section analyzes the cross-section and intertemporal stability of the schooling and experience coefficients estimated in the earnings functions.

Data sources

This section describes the CPS and the decennial census data used to estimate the earnings functions. It discusses the selection of respondents from the CPS and census surveys for inclusion in the sample used to estimate earnings functions. In general, earnings functions are estimated for employees within the private business sector for whom information on earnings and worker characteristics were consistently available from surveys over the period 1950-91.

Description of the CPS and decennial census. Earnings functions are estimated for the years 1949 and 1959 from the decennial census and for 1967-90 from the March Supplementary CPS. Samples were selected from the decennial censuses which include all individual records from the 1 in 1,000 Public Use Sample of Basic Records from the 1960 census and all "sample line" records from the 1 in 100 Censuses of Population, 1950: Public Use Micro Data Sample. All individual records from the CPS are used.

The samples are selected to be consistent with the private business sector of the economy for which the Bureau of Labor Statistics measures labor and multifactor productivity. The BLS study intentionally excludes industrial wage differentials. Therefore, only one set of labor prices is used to calculate earnings weights for both the private business and private nonfarm business sector measures of labor composition. The samples exclude all persons identified as government workers (including postal workers) and persons working in the private household sector. Clergy, religious workers, and persons working in nonprofit membership organizations are also excluded on the assumption that virtually all these persons work in the nonprofit sector. There was not enough information to identify other employees of nonprofit institutions such as those working in hospitals or private schools, so they remain in the sample. The inclusion of some nonprofit sector workers presumably has little effect on the estimated coefficients provided that wages of comparable workers in the profit and nonprofit sectors are similar.

Self-employed persons are also excluded from the sample. Proprietors' income includes returns to both labor and physical capital. It is not possible to separate these two types of income without making debatable assumptions. Nor is it valid to treat all self-employment income as earnings because it would incorrectly ascribe the payments to both capital and labor to just the latter. In effect, the BLS labor input index assigns the hourly price of labor of a wage earner to a self-employed person with the same worker characteristics.

Persons who reported no schooling are excluded from the sample on the assumption that they are either not responding accurately or they received their education outside the traditional classroom. Earnings of these persons are frequently higher than those with 1-4 years of schooling.

Hourly earnings, rather than weekly or annual earnings, are used as the dependent variable in the earnings function to be consistent with the hours series used to construct the BLS multifactor productivity measures. Although not reported directly, hourly earnings can be constructed from data available in the March CPS. For the survey years 1968 through 1975, hourly earnings are calculated from information reported on ment income are excluded. Persons working in private households, general government, clergy, religious workers, and employees of nonprofit membership organizations are removed so that the estimated price of labor corresponds to the sectors for which BLS measures labor and multifactor productivity.

The coverage of hours described in appendix D is more comprehensive than that used for estimating the earnings functions of the private business sector workers. Only those persons who failed to report the number of weeks worked per year or hours worked per week are excluded from the hours measures. Proprietors, unpaid family workers, and persons reporting no education are included in the matrix of hours worked. These persons are assigned an hourly price of labor based on their experience and education as determined by the estimated earnings functions. Persons with no education are assigned an hourly price of labor equal to that of persons with 1-4 years of schooling and other similar worker characteristics.[6]

Data problems. Several problems arise from the manner in which the census and CPS report some of the data. First, extremely high incomes are truncated to protect the confidentiality of the respondents. Second, the number of weeks worked in a given year is reported prior to the 1976 CPS in intervals only. Similarly, the number of hours worked per week in the 1960 decennial census is reported in intervals only. Third, the usual number of hours worked per week in the previous year is only available beginning with the 1976 CPS. Prior to that date, only the number of hours worked in the survey week is reported. Fourth, the method of imputing incomes of nonrespondents was changed in 1988. Finally, answers to certain useful survey questions are not collected in the CPS on a continuous basis.

Annual earnings are truncated at different levels, depending on the year of the survey. Since high earnings are associated with high levels of education and work experience, the truncation reduces the returns and the price of labor associated with the highest levels of education and experience. In addition, the earnings profiles by both education and work experience become flatter in later years as the secular increase in earnings truncates an ever increasing proportion of the sample.[7] Because the proportion of the employed work force in the highest education groups has increased over time, the use of truncated earnings would result in a downward bias in the contribution of education to the growth of labor input.

For each year, a mean estimate of wage and salary income for those persons above the truncation point is substituted for the maximum reportable earnings. The estimate is made by comparing the CPS earnings distribution published in the P-60 series (on money income) to the computer coded sample distribution which uses a much lower truncation point.[8]

The wage bill for all persons below the truncation point is the product of their mean annual earnings and the published number of such wage earners. This wage bill is subtracted from the P-60 published wage bill for all persons to yield an estimate of the total wages and salaries of persons above the truncation point. Mean earnings above the truncation point is obtained by dividing by the published number of high wage earners. The estimated average earnings for those above the truncation point is contained in table E-1.[9]

The second problem is that the number of weeks worked in the previous year is reported directly beginning in 1976; prior to that survey, the data are grouped. For example, persons working between 40 and 47 weeks per year are all counted within the same reported code. To adjust for this, the 1975 mean for each group is obtained from the 1976 March Supplemental CPS, and these are applied to the data for 1959 and 1967-74.[10]

Table E-1. Estimated mean earnings for persons with censored reported wages, 1950-91

Survey year	Truncation point	Estimated mean[1]
1950[2]	$10,000	$13,989
1960[2]	25,000	40,913
1968	50,000	85,000
1969	50,000	83,000
1970	50,000	81,103
1971	50,000	76,078
1972	50,000	80,155
1973	50,000	82,737
1974	50,000	66,414
1975	50,000	76,352
1976	50,000	74,350
1977	50,000	72,325
1978	50,000	70,300
1979	50,000	68,227

[6] The earnings of persons with no education most closely resemble those of workers with 1-4 years of schooling. As a result, persons without education are assigned the same weight. It would have been possible to include all these workers in one group, 0-4 years of schooling, but the result would have been a higher earnings weight for all these workers. If persons reporting no education are simply misreporting, pooling these two groups would overstate the earnings of all. Instead, if these persons truly have no education, but do have higher earnings, omission of these workers results in a sample selection bias in the education coefficients.

[7] Truncation lowers estimated hourly earnings of white male postgraduates by less than 1 percent in the 1968 CPS, but by more than 8 percent in 1980. Earnings are truncated at $50,000 in both years.

[8] The published wage data come from the P-60 series of the CPS on consumer income. The exact title differs slightly from year to year, but the 1980 data, for example, can be found in table 53 titled Type of Income, by Income of Specified Type in 1980—Persons 15 years Old and Over, by Race, Spanish Origin, and Sex, in Series P-60 No. 132, Money Income of Households, Families, and Persons in the United States: 1980.

[9] This method is unavailable for 1950 and 1960. The truncation point in those years is $10,000 and $25,000 respectively. Using the March 1968 CPS, the mean earnings above $10,000 and $25,000 were substituted.

[10] The following 1975 means are substituted for weeks worked during the prior year:

Group weeks	Mean weeks
1-13	8.72
14-26	21.75
27-39	34.24
40-47	43.31
48-49	48.30
50-52	51.89

Table E-1. Estimated mean earnings for persons with censored reported wages, 1950-91–Continued

Survey year	Truncation point	Estimated mean[1]
1980	50,000	73,810
1981	50,000	67,835
1982	75,000	95,967
1983	75,000	94,482
1984	75,000	93,821
1985	99,999	137,003[3]
1986	99,999	137,003
1987	99,999	140,026
1988	99,999[4]	144,617
1989	99,999[4]	144,986
1990	99,999[4]	156,053
1991	99,999[4]	145,467

[1] Mean earnings for persons with earnings above the truncation point.

[2] The 1950 and 1960 figures are based on the March 1968 CPS earnings of those above $10,000 and $25,000, respectively.

[3] In 1985, both the P-60 published series and computer accessible earnings are truncated at the same level. The 1986 estimated mean is substituted.

[4] Prior to 1988 the truncation point was for all earnings. After 1988, earnings from each of the three longest jobs was reported and subject to truncation. Some workers may indicate earnings from all three jobs that are larger than the estimated mean. The higher value is used.

The same problem occurs in reporting the number of hours worked per week in the 1960 decennial census. The March 1971 CPS is used to construct point estimates for 1959.[11]

The third major difficulty arises from the change in the measure of hours worked per week. Prior to the 1976 survey, the sole measure of weekly hours is the reported number of hours worked in the survey week, whereas earnings and weeks worked per year relate to the previous year.[12] As a result, the hourly earnings function does not, strictly speaking, yield a measure of labor prices for either the current or previous year. Because average weekly hours tend to lead the business cycle, the measure of average hourly earnings may be somewhat overstated during years of recession and understated during years of recovery. Furthermore, the use of average hours worked in the survey week may result in slight differences in the distribution of earnings from the actual distribution in the year prior to the survey. Given that the annual earnings and weeks worked data apply to the year prior to the survey, an hourly wage rate based on annual earnings, weeks per year, and hours per week is most consistent with the year prior to the survey year.

Beginning with the 1988 survey, the CPS uses a new method to impute the income of nonrespondents. Since the imputation is based on the known characteristics of the respondent, the estimated parameters of the wage model will be affected. Changes between the 1987 and 1988 surveys in the estimated parameters will then reflect changes in the reward to characteristics as well as changes in the imputation method.

The 1988 survey data contains income measures based on both the old and new imputation methods, and two wage equations are estimated. A wage equation using data based only on the old imputation method is consistent with wage equations estimated prior to the 1988 survey. A second wage equation using data based only on the new imputation method is consistent with wage equations for surveys after 1988. By using two sets of 1988 estimated parameters to link the wage equations, changes in parameter estimates due to changes in the method of imputation are eliminated.[13]

Another, but less significant, problem occurs because the sample was altered again in the 1980 survey when the educational attainment of 14-year-olds was no longer reported. Beginning in 1979, the wage equations do not include 14-year-olds in the sample.

Still another issue arises from the fact that certain information on worker characteristics is not collected on a regular basis in the CPS. Omitting these variables can result in biased estimates of the remaining coefficients, but sporadically including them could yield large and meaningless fluctuations in the estimated annual coefficients. Experiments indicate that the estimated coefficients for schooling and experience, the coefficients of concern here, are fairly robust with respect to the omitted variables problem. (See appendix B for tests of selected omitted variables using the NLS data.) The estimated earnings functions are based on the set of variables which are, with few exceptions, available in all years.

There are also three differences in definition between the CPS and decennial census data that can cause some problems of comparability. Respondents to the CPS are identified as part-time workers provided they work 1-34 hours per week for a majority of the weeks in which they worked the previous year. This differs conceptually from the definition used in the 1950 and 1960 decennial censuses. In these censuses, a part-time worker is anyone working 1-34 hours in the week prior to enumeration. This includes persons voluntarily (such as for illness or vacation) or involuntarily (such as for slack work or material shortages) on part-time schedules in the survey week. Their annual salary, however, reflects their usual full-time work schedule, and the lower estimate of hours worked inflates their hourly earnings. The hourly wage rate of self-identified part-time workers is expected to be lower than the wage rate of persons working 1-34 hours in a given week; this is exactly what was found. Since self-identified part-time

[11] The following 1971 means are substituted for hours worked in the survey week:

Group hours	Mean hours
1-14	8.28
15-29	20.95
30-34	31.39
35-39	36.43
40	40.00
41-48	45.57
49-59	52.04
60 or more	68.89

[12] During the survey years prior to 1976, the CPS reported grouped data on average weekly hours for the previous year, but these were not used because it was believed that the point estimates calculated from the data would likely yield larger errors than the approach described in the text.

[13] Reported parameter estimates are based on data using the old imputation method. The other set of estimates is not reported.

worker status is unavailable from the decennial censuses, no variable defining part-time worker status is included in the earnings functions for those particular years.

The decennial censuses report the number of children ever born to women, but the Current Population Survey reports only the number of children of any age and marital status living in the household or away attending school. This definitional difference affects the measures of predicted work experience used in the wage equations. For younger women, the two measures are virtually identical. However, for relatively older women, these children may be adults and no longer living at home. In an identical manner to the experience equations described in appendix C, the proportion of older women with zero, 1, 2 or 3, or 4 or more children from the 1 in 100 sample of the 1970 decennial census is substituted for the CPS data for women who completed their schooling at least 26 years ago.[14] Five separate sets of proportions classified by education are used for ever married women. Another set of proportions are used for never married women.

One minor discrepancy remains. Residence within a standard metropolitan statistical area (SMSA) is divided into central city (CCITY) and balance of the city (BCITY) in the CPS. No division of residence within a SMSA is made in the decennial census. Instead, residence is either within a SMSA (URBAN) or not. Differences in wages between residents of the central city and residents of the balance of the city are quite small for men as well as for women prior to 1980. Therefore, the abbreviated classification available in the decennial census is not a serious shortcoming.

Earnings functions

This section describes the specification of the earnings functions in more detail than in appendix A. It also presents the estimated results. As indicated above, earnings functions are estimated separately for men and women. These functions are used to calculate the price of labor for all men and all women regardless of race or ethnicity.

Earnings equation. The earnings function is a semilog form. It is basically an extension of the equation developed by Mincer in that it incorporates other worker characteristics beyond schooling and experience.[15] The earnings functions provide estimates of payments per hour associated with worker characteristics. Thus, hourly earnings are interpreted as the amount firms in the private business sector are willing to pay for the services of a worker with a given set of characteristics, including the human capital investments associated with schooling and work experience.

The variables included in the earnings functions are listed in table E-2. Acronyms for the regressors are shown in parentheses. Dummy variables are used for different schooling groups; the base group includes persons with 5-8 years of completed schooling.

Table E-2. Variables in the earnings functions

Dependent variable:
Natural log of hourly wages
Explanatory variables:
Estimated SSA work experience (EXPER)
Estimated SSA work experience squared (EXPSQ)
Dummy variable for those:
with 1-4 years of completed schooling (S1-4)
with 5-8 years of completed schooling (S5-8) (omitted variable)
with 9-11 years of completed schooling (S9-11)
with 12 years of completed schooling (S12)
with 13-15 years of completed schooling (S13-15)
with 16 years of completed schooling (S16)
with more than 16 years of completed schooling (S17+)
Dummy variable for:
part time employment (PTIME) (available 1968-91)
veterans (VET) (men only prior to 1989)
resident of New England Census Division (NEWENG)
resident of Mid-Atlantic Census Division (MIDATL)
resident of East North Central Census Division (ENCENT)
resident of West North Central Census Division (WNCENT)
resident of South Atlantic Census Division (SATL)
resident of East South Central Census Division (ESCENT)
resident of West South Central Census Division (WSCENT)
resident of Mountain Census Division (MOUNTAIN)
resident of Pacific Census Division (PACIFIC) (omitted variable)
resident of central city of SMSA (CCITY) (available 1968-91)
resident of balance of SMSA (BCITY) (available 1968-91)
resident of SMSA (URBAN) (available 1950 and 1960)
residents living outside SMSA (RURAL) (omitted variable)

For example, the estimated coefficient with respect to the "dummy variable for those with 12 years of completed schooling (S12)" measures the approximate percentage difference between the hourly pay of an employee with 12 years of schooling and an employee with 5-8 years of schooling (that is, the base group), holding experience and other worker

[14] The proportions are based on samples of women who completed their schooling at least 26 years ago and reported a positive number of hours and weeks in the previous year. See appendix C for further details.

[15] See Mincer (1974) for a derivation of empirical wage models. While it is possible to derive estimating forms from assumptions concerning the optimal human capital investment path, this has not been done for this model. Ben-Porath (1967) derived a theoretically consistent model of human capital investment, but it is intractable to estimation. Y. Ben-Porath does not seek to derive estimating forms. Instead, he is concerned with the optimal life cycle path for human capital investment.

characteristics constant. The educational groupings in the earnings functions are the same as those for hours shown in appendix D.

The variable for work experience in the earnings function uses the values obtained from the work experience equation employed in the first stage of the two stage procedure described in appendix C. Linear and quadratic experience terms are included in the earnings function; this implies a parabolic experience-earnings profile consistent with commonly observed age-earnings profiles.

The remaining variables in the earnings function relate earnings to worker and job characteristics other than schooling and work experience. Although they are essentially used as control variables, some explanations may be useful because several can be viewed as related to investment in human capital.

Hourly cash earnings of part-time workers are generally less than earnings of full-time workers. To take account of this, a dummy variable is included for part-time workers in the earnings functions. Several interpretations are possible. Persons working part-time schedules (voluntarily or not) may earn less because they are less willing or have less opportunity to invest in on-the-job training than full-time workers.[16] Or, these workers may be receiving a negative compensating differential for their short workweek. Successful job sorting by employees would lead to the conclusion that workers with the strongest preference for the short workweek will occupy these jobs. These workers are, therefore, the most willing to sacrifice wages to achieve a part-time work schedule.[17]

In addition, the firm's fixed cost per employee may make a part-time schedule more costly per hour of work, and it may be less willing to pay the full-time hourly rate. Finally, in addition to the possibility that part-time workers invest in skills at a slower rate, there may be unobservable differences between part- and full-time workers. The fact that some workers are on part-time schedules may indicate that firms have a low regard for these workers because of unmeasured differences in abilities, motivation, or other worker qualities. In sum, it is unclear whether the hourly cash wage rate of part-time workers is less than that of full-time workers because of differences in the level of skills or preference for a shorter workweek. This study assumes the latter.

A dummy variable for veterans is included in the earnings function for men (and for women beginning in 1989). Several studies have shown, other things equal, that earnings of veterans of World War II and Korea are higher than those of nonveterans, but more recent veterans either receive no premium or have lower earnings.[18] The earnings differential for veterans may reflect direct effects of military training or service, unmeasured quality differences, the effect of veteran status as a screen for either quality differences or subsequent training, and differential treatment of veterans in the labor market.[19] The measures of schooling and experience would not pick up this earnings difference because the rate of investment in skills is assumed to be the same in both civilian and military activity.[20] One could possibly make a case for including the variable for veterans in the calculation of the price of labor, but it was not done because it would require increasing the dimension of the hours matrices, and it is not clear, in any case, what the variable actually measures.

The earnings functions are estimated without regard to race or ethnicity. The samples of white, black, and all ethnic groups are pooled to estimate a common set of coefficients. As indicated in appendix A, racial and ethnic discrimination is assumed to take the form of employee discrimination. In this case, the earnings of neither whites nor minorities equal their marginal product of labor. Since employee discrimination primarily redistributes earnings between workers, the average wage paid to all workers still measures the marginal product of labor for all workers.[21]

Finally, dummy variables for census divisions and urban localities are included in the earnings functions. These are intended to reflect wage differentials induced by regional price differences relative to the omitted group, nonurban residents of the Pacific census division.

[16] The rate of training for part-time workers is examined by Corcoran, Duncan, and Ponza (1983) and Duncan and Hoffman (1979).

[17] From the perspective of individual utility maximization, a person determines the optimal number of hours to work at the prevailing wage. Any deviation from this optimum is an inferior solution. Therefore, a person wishing to work only part time can be made indifferent to working full time only by some form of compensating differential. A full-time worker also requires a compensating differential to be induced to work part time. Profit maximizing firms can minimize labor costs by placing workers in their most desired work schedule.

If all persons work their optimal number of hours then no measurable compensating differential will be apparent. A negative differential for part-time workers suggests that there are an insufficient number of part-time jobs relative to demand, and a positive differential suggests the opposite.

In Brown's survey (1980) of 11 studies of compensating differentials, he notes that empirical studies frequently fail to observe compensating differentials or estimates may conflict with theory. This seems to be the case in his survey of estimates for job safety and working conditions. However, the few studies that examine work schedules all find lower hourly earnings for those with flexible hours. It would appear that this premium paid by employees who can set their work schedule implies a desire for a flexible work schedule. Also consistent with the notion that 40-hour workweeks are in oversupply, Brown finds that part-time workers are paid a significantly lower hourly wage than full-time workers. Hourly wages are also significantly lower for persons working more than 40 hours per week compared to persons working just 40 hours.

[18] Berger and Hirsch (1983) find that Viet Nam Era veterans earn almost as much as nonveterans. However, they find that Korean, World War II, and other veterans earn significantly more than nonveterans.

[19] See Goldberg and Warner (1987) for a discussion of this point.

[20] Military service is generally credited toward Social Security coverage and is therefore included in the measures of work experience used in this study. Similarly, men sent to college during their military service would also be credited with this schooling. However, training at technical schools would not be included.

The most important issue concerns the intensity of training and schooling while in military service. If the intensity differs from civilian work, veterans acquire skills at a different rate than civilians, and the veteran status dummy variable reflects skill differences. This study assumes veterans acquire skills at the same rate as civilians.

[21] Employee discrimination also distorts the allocation of resources. This implies that the value of marginal products for some type of capital and labor are not equal to their factor prices. Therefore, it is possible that the average wage may differ from the marginal product of all workers.

The price of labor used to weight the hours of work is not identical with earnings derived directly from the earnings function. A simplified version of the earnings function presented in table E-2 is shown as equation 1.

Earnings function:

$$1) \quad \text{Ln } W = a + b * ED + c * EX - d * EX^2 + f * Z$$

Specifically, the price of labor, W, is the sum of the payments for education, ED, work experience, EX, and unmeasured skills, $a + f * Z$. The price of labor W′ in equation 2 adds to the intercept the mean value of the other traits, $\overline{Z}$, weighted by the estimated coefficients f, that is $a' = a + f * \overline{Z}$.[22]

The price of labor:

$$2) \quad \text{Ln } W' = a' + b * ED + c * EX - d * EX^2$$

The excluded traits, Z, may or may not represent productivity-related differences between workers. Regardless, firms are assumed to equate wages with the value of marginal products, and so these differences are relevant to the measurement of output elasticities. Since these other traits are numerous, inclusion would partition the hours matrix too thinly to measure meaningfully the hours of each type of worker. Instead, the values of these other traits are included in a modified intercept. As indicated in appendix A, education and experience parameters which include the return to education and experience as well the return to other traits yield an ambiguous measure of labor composition. To minimize this problem, these other traits are included in the earnings function so that the measures of returns to education and experience more closely reflect those returns and no other extraneous factors. By adding a weighted average of these other traits to the intercept, the adjusted intercept approximates the returns to skills other than education and experience.[23]

Differences between men and women in the intercept are permitted to remain. As indicated in appendix A, a number of interpretations are possible. For the purposes of measuring labor composition, two interpretations deserve concern: Differences in unmeasured skills and discrimination. By adjusting separately the intercepts for men and women, in effect, the regional and urban dummy variables for men and women have been equalized.[24] The difference between the male and female intercepts will now reflect the U.S. average level of sex discrimination, if any, as well as differences in unmeasured skills.

Estimated earnings function. As indicated earlier, the hourly prices of labor used in the labor input index are calculated from the schooling and experience coefficients estimated in the earnings functions. These estimated coefficients (and their t-statistics) for men and women based on the 1950, 1960, and 1968-91 surveys are shown in tables E-3 and E-4. Tables E-12 and E-13 at the end of the appendix show, for selected years, the complete estimated earnings functions described in table E-2.

In the case of men, the estimated schooling coefficients are all significantly different from zero (all tests are at the 95-percent confidence level); and, with only two exceptions (S17+ in 1950 and 1971) the coefficients in each year increase with additional years of schooling, as expected. Similarly, all of the estimated coefficients with respect to the experience variable have the expected signs and are statistically significant.

It is more difficult to appraise the stability, or lack of stability, of the estimated coefficients over time because changes can reflect "real" economically induced short-run and long-run movements as well as random and/or systematic measurement errors. For men, experience coefficients are more stable than those for schooling. But, on the whole, even the annual changes in the estimated schooling coefficients do

[22] For each year, equation 2 weights the mean value of all variables except education and work experience by their estimated coefficient and adds this to the estimated intercept.

[23] In effect, this study does not consider the effect of differences in the distribution of these traits by education and experience. However, differences in the distribution of these traits between men and women are captured by the use of separate equations for men and women.

[24] The adjustment to the intercept is consistently small and ranges in value from -0.031 to 0.040. These adjustments change the wage of women as well as the male/female relative earnings by no more than 4 percent. (These adjustments usually cause a change of less than 10 percent in the estimated female intercept, but the change does range from -11.2 percent to 45.9 percent.)

Table E-3. Estimated earnings function coefficients for men, selected years, 1950-91

Survey year	Intercept	Years of education						Experience[1]	
		1-4	9-11	12	13-15	16	17+	Linear	Quadratic
1950	-0.354	-0.150	0.093	0.163	0.245	0.422	0.447	0.015	-0.000078
	(23)	(12)	(11)	(19)	(19)	(25)	(18)	(48)	(37)
1960	.133	-.156	.100	.179	.324	.545	.550	.015	-.000074
	(7.5)	(8.3)	(10)	(18)	(25)	(33)	(24)	(40)	(29)
1968	.345	-.199	.102	.231	.351	.520	.627	.014	-.000070
	(16)	(7.7)	(7.9)	(19)	(23)	(29)	(25)	(28)	(21)
1969	.344	-.132	.112	.242	.364	.531	.638	.015	-.000080
	(16)	(4.9)	(8.6)	(20)	(24)	(29)	(26)	(31)	(25)
1970	.508	-.205	.137	.243	.349	.554	.645	.014	-.000071
	(24)	(7.6)	(10)	(20)	(24)	(32)	(27)	(28)	(22)
1971	.475	-.194	.112	.241	.395	.572	.553	.015	-.000082
	(21)	(6.7)	(8.0)	(19)	(25)	(31)	(24)	(30)	(25)
1972	.425	-.198	.130	.239	.362	.541	.652	.018	-.000098
	(19)	(6.7)	(9.1)	(18)	(23)	(29)	(27)	(38)	(30)
1973	.531	-.234	.088	.209	.313	.499	.568	.017	-.000088
	(23)	(7.5)	(5.8)	(15)	(19)	(26)	(23)	(34)	(27)
1974	.540	-.237	.155	.250	.326	.485	.615	.018	-.000096
	(23)	(7.2)	(9.9)	(18)	(20)	(25)	(26)	(36)	(28)
1975	.671	-.268	.124	.214	.275	.431	.533	.018	-.000095
	(26)	(7.1)	(7.1)	(13)	(15)	(21)	(21)	(33)	(26)
1976	.584	-.298	.117	.259	.336	.554	.662	.019	-.000105
	(27)	(9.1)	(7.8)	(19)	(22)	(32)	(31)	(45)	(36)
1977	.644	-.158	.117	.275	.337	.562	.686	.019	-.000103
	(33)	(5.2)	(8.6)	(22)	(24)	(36)	(36)	(50)	(39)
1978	.684	-.208	.154	.305	.376	.571	.695	.020	-.000104
	(34)	(6.7)	(11)	(24)	(26)	(35)	(36)	(49)	(38)
1979	.805	-.235	.162	.317	.392	.589	.706	.018	-.000093
	(39)	(7.0)	(11)	(23)	(26)	(35)	(36)	(45)	(34)
1980	.921	-.273	.132	.294	.364	.545	.699	.019	-.000099
	(47)	(8.5)	(8.9)	(22)	(25)	(34)	(38)	(49)	(38)
1981	1.016	-.204	.142	.293	.351	.547	.704	.019	-.000098
	(51)	(6.3)	(9.4)	(22)	(24)	(34)	(38)	(49)	(38)
1982	1.038	-.267	.140	.281	.358	.539	.686	.019	-.000097
	(45)	(6.9)	(8.0)	(18)	(21)	(30)	(34)	(44)	(33)
1983	1.042	-.239	.124	.305	.371	.562	.735	.020	-.000102
	(42)	(5.7)	(6.3)	(17)	(20)	(28)	(34)	(43)	(33)
1984	1.023	-.183	.138	.294	.379	.598	.782	.020	-.000105
	(40)	(3.9)	(6.8)	(16)	(20)	(29)	(35)	(45)	(34)
1985	1.079	-.210	.130	.313	.377	.647	.777	.020	-.000101
	(44)	(5.0)	(6.7)	(18)	(20)	(33)	(36)	(45)	(34)
1986	1.047	-.179	.183	.329	.447	.664	.870	.020	-.000102
	(42)	(4.1)	(9.1)	(18)	(23)	(33)	(39)	(45)	(33)
1987	1.086	-.215	.143	.318	.421	.697	.865	.020	-.000102
	(44)	(4.9)	(7.2)	(18)	(22)	(35)	(39)	(45)	(33)
1988	1.115	-.041	.166	.335	.443	.694	.881	.019	-.000091
	(51)	(1.1)	(9.4)	(21)	(26)	(39)	(45)	(47)	(33)
1989	1.143	-.108	.185	.349	.462	.721	.882	.018	-.000089
	(49)	(2.8)	(9.8)	(20)	(26)	(38)	(42)	(43)	(30)
1990	1.248	-.097	.144	.328	.474	.738	.900	.017	-.000080
	(57)	(2.7)	(7.8)	(20)	(27)	(40)	(44)	(41)	(28)
1991	1.276	-.111	.172	.317	.485	.745	.899	.017	-.000084
	(57)	(2.9)	(9.0)	(19)	(27)	(40)	(43)	(42)	(30)

[1] Experience is measured in quarters of estimated work experience.

NOTE: t-statistics in parentheses.

Table E-4. Estimated earnings function coefficients for women, selected years, 1950-91

Survey year	Intercept	Years of education						Experience[1]	
		1-4	9-11	12	13-15	16	17+	Linear	Quadratic
1950	-0.363	-0.015	0.072	0.145	0.247	0.297	0.313	0.0075	-0.000049
	(15)	(.5)	(5.2)	(12)	(13)	(10)	(7.2)	(13)	(8.7)
1960	.048	-.117	.070	.163	.266	.403	.453	.0073	-.000040
	(1.7)	(2.9)	(4.3)	(11)	(13)	(12)	(8.5)	(11)	(6.5)
1968	.087	-.064	.085	.228	.323	.493	.574	.0110	-.000081
	(2.6)	(1.0)	(3.9)	(12)	(13)	(13)	(9.1)	(14)	(10)
1969	.277	-.121	.070	.205	.329	.427	.632	.0088	-.000060
	(8.5)	(2.1)	(3.2)	(11)	(13)	(12)	(11)	(11)	(7.9)
1970	.402	-.099	.049	.170	.302	.380	.376	.0082	-.000057
	(13)	(1.7)	(2.4)	(9.4)	(13)	(12)	(7.6)	(12)	(8.4)
1971	.341	-.108	.102	.252	.390	.504	.359	.0102	-.000074
	(11)	(1.8)	(4.8)	(13)	(17)	(15)	(7.9)	(14)	(10)
1972	.328	-.085	.113	.224	.322	.476	.637	.0115	-.000083
	(10)	(1.4)	(5.0)	(11)	(13)	(14)	(12)	(15)	(11)
1973	.405	-.069	.095	.196	.303	.428	.451	.0117	-.000088
	(12)	(1.2)	(4.1)	(9.5)	(12)	(13)	(8.9)	(16)	(12)
1974	.452	-.063	.100	.192	.340	.416	.532	.0120	-.000082
	(14)	(1.0)	(4.2)	(8.9)	(14)	(13)	(11)	(16)	(11)
1975	.566	.020	.055	.164	.292	.378	.508	.0114	-.000081
	(17)	(.3)	(2.2)	(7.3)	(11)	(12)	(11)	(14)	(10)
1976	.547	-.081	.052	.216	.331	.402	.462	.0128	-.000097
	(20)	(1.4)	(2.5)	(12)	(16)	(16)	(13)	(20)	(15)
1977	.587	-.061	.067	.214	.326	.456	.576	.0139	-.000107
	(24)	(1.2)	(3.7)	(13)	(18)	(20)	(18)	(25)	(19)
1978	.683	-.049	.094	.245	.360	.487	.659	.0123	-.000086
	(30)	(1.0)	(5.2)	(15)	(20)	(23)	(22)	(23)	(16)
1979	.840	-.018	.042	.216	.316	.404	.552	.0104	-.000070
	(35)	(.3)	(2.1)	(12)	(16)	(18)	(19)	(19)	(12)
1980	.891	-.070	.093	.230	.340	.507	.609	.0122	-.000091
	(37)	(1.3)	(4.8)	(13)	(18)	(23)	(22)	(22)	(16)
1981	.996	.054	.081	.219	.350	.501	.559	.0125	-.000093
	(41)	(1.0)	(4.1)	(12)	(18)	(23)	(20)	(23)	(17)
1982	1.093	-.175	.036	.200	.343	.453	.619	.0116	-.000083
	(43)	(3.1)	(1.7)	(10)	(17)	(20)	(23)	(20)	(14)
1983	1.008	-.222	.096	.244	.370	.534	.650	.0161	-.000130
	(36)	(3.5)	(4.1)	(12)	(16)	(22)	(22)	(26)	(20)
1984	1.063	-.217	.080	.246	.395	.539	.712	.0140	-.000099
	(35)	(3.2)	(3.1)	(11)	(16)	(20)	(23)	(21)	(14)
1985	1.071	.002	.085	.236	.402	.561	.734	.0162	-.000120
	(37)	(.0)	(3.5)	(11)	(17)	(22)	(25)	(26)	(18)
1986	1.091	-.120	.112	.258	.410	.617	.768	.0161	-.000119
	(38)	(1.7)	(4.6)	(12)	(18)	(25)	(26)	(26)	(18)
1987	1.064	.015	.063	.250	.404	.624	.806	.0167	-.000124
	(35)	(.2)	(2.4)	(10)	(16)	(24)	(26)	(24)	(17)
1988	1.030	-.031	.161	.323	.495	.681	.856	.0176	-.000131
	(39)	(.6)	(7.0)	(15)	(23)	(30)	(32)	(22)	(21)
1989	1.148	-.117	.128	.296	.482	.715	.859	.0148	-.000107
	(41)	(2.2)	(5.2)	(13)	(21)	(29)	(31)	(25)	(18)
1990	1.191	-.073	.100	.284	.454	.726	.797	.0162	-.000121
	(43)	(1.2)	(4.1)	(13)	(20)	(30)	(29)	(27)	(20)
1991	1.210	.006	.137	.310	.481	.739	.853	.0167	-.000127
	(43)	(.1)	(5.5)	(14)	(20)	(30)	(31)	(28)	(21)

[1] Experience is measured in quarters of estimated work experience.

NOTE: t-statistics in parentheses.

not appear to be highly volatile except for the lowest education group (1-4 years) for which the number of observations is comparatively small.

The results for women are not quite as consistent as those for men, in part, because the sample is smaller for women. There are seven cases for which the estimated schooling coefficients do not change as expected with the years of schooling, but these are for the lowest and highest education groups and, in part, reflect the small number of observations in these cells. All of the estimated experience coefficients, on the other hand, have the expected signs and are statistically significant. The schooling and experience coefficients are more volatile for women than for men.

Tables E-5 and E-6 show the indexes of the hourly price of labor by years of schooling, holding experience constant.[25] These are calculated from the estimated coefficients reported in tables E-3 and E-4. The figures for men in 1991, for example, indicate that the cost of an hour's work was 135.4 (235.4-100) percent higher for a person with 16 years of schooling than for a person with 1-4 years of schooling, and the cost was 53.4 (235.4/153.5-100) percent higher than that of a high school graduate. On the whole, price differentials are larger for men than for women. For example, the difference in the price of labor for an hour's work for a person with a college degree and one with a high school diploma is larger for men than for women. The comments made above concerning tables E-3 and E-4 obviously apply here also: The price for an hour's labor always increases with years of schooling with two exceptions for men and in nearly all cases for women.

The earnings weights actually used to compute the labor input index are more stable than the annual prices of labor shown in the tables. This is because the prices of labor used in the Tornqvist index number formula are averages for two consecutive years (see appendix A). Use of these averages tends to reduce the effects of any random fluctuations.

Tables E-7 and E-8 show indexes of the relationship between the hourly price of labor and years of experience, holding education constant. These, too, are calculated from the estimated coefficients shown in tables E-3 and E-4. Besides these hourly labor price-experience profiles, the tables show the peak earnings point, that is, the number of years of experience at which earnings reach their peak. The tables also show the marginal percentage increase in the price of labor resulting from an additional year of work experience (evaluated at the annual sample mean of experience). In the case of men, the labor price-experience profiles have tilted upward over the 41-year period, 1950 through 1991. Thus, in 1991, an hour's work by a man with 20 years of experience was 134.1 percent higher than that of a man with no experience, whereas in 1950, the differential was 86.9 percent. The peak earnings year for men has remained fairly constant, about 24 years, during the entire period. The marginal increase in the price of labor associated with an additional year of experience rose substantially over the period reflecting the upward tilt in the profiles and the decline in the mean level of work experience.

The picture is much the same for women: The hourly labor price-experience profile tilted sharply upward during the 41-year period to the degree that the marginal rate of increase in the price of labor nearly doubled. The peak year may have decreased slightly during the period, but this is hard to discern because of the annual fluctuations in the series. However, the flatter price-experience profile for women is the most striking difference between the price of labor for men and women.

Additional statistical tests

This section presents the results of two additional econometric tests related to the earnings functions. The results of the first test show that the earnings functions for men and women differ significantly, and that these differences persisted over the 41-year period covered in this bulletin. The second test shows that the annual movements in the estimated education and experience coefficients shown in tables E-3 and E-4 reflect significant intertemporal changes in the earnings functions, particularly with respect to the labor price-experience profiles.

Tests for differences by sex. There are a host of studies of male-female wage differentials based on earnings functions which have found that the estimated coefficients for education and experience are significantly different for men and women. These studies have been based on many different sources of data covering a variety of different time periods.[26] The results presented here, based on decennial census and CPS data, are consistent with the conclusions reached in the earlier studies.

To test for significant differences between the estimated coefficients for men and women, the samples for both genders are first pooled. An earnings function is then estimated from these pooled data which includes the variables specified in table E-2 and interaction terms between women and years of schooling, experience, and the intercept.

The use of interaction terms permits differences between men and women to be measured and tested. These functions are estimated for 6 selected years covering the period beginning in 1950 and ending in 1983. The top half of table E-9 shows the estimated coefficients (and t scores) for men. The returns to women are measured by the sum of the coefficients on these variables and the corresponding interaction term. The difference in the returns of men and women are measured by the interaction terms and are shown in the bottom half of the table (along with t scores). For example, the estimated coefficient for the return to male college graduates in 1950 is 0.432. The corresponding return for women is 0.273 (= -0.159 + 0.432). The interaction term between college graduates and the female dummy variable has a coefficient of -0.159 implying that female college graduates earn less than male college graduates in that year. This difference was also significantly different from zero. More generally, the results indicate that

[25] The semilog form of the earnings function without interaction terms between education and experience yields an index of the price of labor by educational attainment that is identical at all levels of experience.

[26] For a survey and bibliography to earlier studies, see Francine D. Blau and Marianne A. Ferber (1986). Also, see Solomon W. Polacheck (1984).

Table E-5. Indexes of the hourly price of labor by years of education for men, 1950-91
(1-4 years of education = 100)

Survey Year	Years of schooling					
	5-8	9-11	12	13-15	16	17+
1950	118.1	131.0	139.7	152.8	182.1	177.9
1960	116.8	129.2	139.7	161.5	201.6	202.6
1968	122.0	135.1	153.8	173.4	205.2	228.4
1969	114.1	127.7	145.4	164.3	194.0	216.0
1970	122.8	140.8	156.6	174.1	213.6	234.0
1971	121.4	135.8	154.5	180.2	215.1	210.9
1972	121.9	138.7	154.7	175.0	209.3	233.8
1973	126.4	138.0	155.7	172.9	208.1	223.1
1974	126.7	149.9	162.7	175.4	205.8	234.3
1975	130.8	148.0	161.9	172.1	201.2	222.7
1976	134.7	151.4	174.6	188.5	234.3	261.0
1977	117.1	131.7	154.1	163.9	205.4	232.4
1978	123.2	143.7	167.2	179.5	218.1	246.9
1979	126.5	148.7	173.7	187.3	228.0	256.3
1980	131.4	149.9	176.3	189.1	226.6	264.2
1981	122.7	141.4	164.4	174.3	212.0	248.0
1982	130.7	150.2	173.0	186.8	224.0	259.4
1983	127.0	143.7	172.2	184.1	222.7	264.8
1984	120.0	137.8	161.0	175.3	218.3	262.4
1985	123.4	140.5	168.8	180.0	235.8	268.5
1986	119.6	143.6	166.2	186.9	232.3	285.5
1987	124.0	143.1	170.5	188.9	249.0	294.5
1988	104.2	123.1	145.7	162.2	208.6	251.6
1989	111.4	134.0	157.9	176.9	229.2	269.1
1990	110.2	127.2	152.9	176.9	230.3	270.9
1991	111.8	132.7	153.5	181.5	235.4	274.5
Mean	121.6	139.1	159.9	176.2	216.8	245.5

Source: Calculated from table E-3.

Table E-6. Indexes of the hourly price of labor by years of education for women, 1950-91
(1-4 years of education = 100)

Survey Year	Years of schooling					
	5-8	9-11	12	13-15	16	17+
1950	102.5	111.1	121.9	131.9	135.0	138.0
1960	112.8	123.1	135.1	149.9	171.7	180.2
1968	106.9	117.0	134.2	147.0	173.9	188.8
1969	111.5	120.8	136.7	154.1	171.4	208.5
1970	111.0	116.5	130.7	148.5	160.8	160.4
1971	111.8	123.6	140.6	160.3	178.7	154.1
1972	109.2	122.8	137.0	150.9	176.0	206.4
1973	107.3	118.6	131.0	145.8	165.4	169.1
1974	107.4	119.2	130.3	151.1	163.0	163.0
1975	98.6	104.4	116.2	131.9	143.8	163.7
1976	108.6	114.9	135.2	151.4	163.1	172.5
1977	106.6	114.7	132.5	148.0	168.4	189.6
1978	105.4	116.4	135.0	151.1	171.6	203.3
1979	102.6	107.3	127.3	140.2	153.2	177.2
1980	108.2	119.2	136.3	151.7	179.4	198.4
1981	95.3	103.9	118.8	135.2	157.4	166.5
1982	120.0	124.9	146.7	168.9	188.5	222.1
1983	125.5	139.0	160.2	181.2	213.2	238.8
1984	125.2	136.4	160.2	185.4	214.3	253.9
1985	100.9	110.4	127.8	150.3	176.3	209.0
1986	114.1	128.3	148.0	171.8	211.4	245.2
1987	99.1	106.2	127.7	148.7	185.4	221.7
1988	103.2	121.3	142.5	169.2	203.8	242.8
1989	112.4	127.7	151.1	182.0	229.6	265.4
1990	107.6	118.9	142.9	169.4	222.3	238.7
1991	99.4	114.0	135.5	160.7	208.1	233.2
Mean	108.2	118.5	136.2	155.3	180.2	201.2

Source: Calculated from table E-4.

Table E-7. Indexes of the hourly price of labor by years of work experience, peak earnings years, and marginal increase in hourly labor prices for men, 1950-91

Survey year	Years of work experience					Peak[1] earnings	Marginal[2] increase in hourly labor price (percent)
	0	5	10	15	20		
1950	100.0	126.9	152.6	186.9	184.0	24.3	1.53
1960	100.0	130.4	160.1	201.9	200.7	24.7	1.66
1968	100.0	127.7	154.1	189.9	187.1	24.3	1.53
1969	100.0	131.2	161.4	201.5	194.7	23.7	1.62
1970	100.0	127.3	153.1	186.7	181.1	23.7	1.44
1971	100.0	130.5	159.6	196.3	185.9	22.9	1.47
1972	100.0	137.9	175.7	225.6	211.5	22.9	1.85
1973	100.0	134.7	169.1	215.6	207.2	23.6	2.00
1974	100.0	137.6	175.3	226.3	215.2	23.4	2.15
1975	100.0	137.0	174.0	223.4	211.8	23.2	2.03
1976	100.0	141.3	183.6	240.9	225.8	23.1	2.24
1977	100.0	141.6	184.7	245.5	235.1	23.7	2.55
1978	100.0	141.8	185.0	245.1	232.6	23.4	2.56
1979	100.0	138.5	178.0	235.6	231.8	24.5	2.60
1980	100.0	140.0	181.1	238.7	228.9	23.7	2.57
1981	100.0	140.2	181.6	240.6	232.6	23.9	2.60
1982	100.0	140.3	182.1	243.2	238.2	24.3	2.61
1983	100.0	142.1	186.2	250.3	242.9	24.1	2.62
1984	100.0	144.1	190.8	260.3	253.9	24.3	2.81
1985	100.0	143.7	190.6	263.2	263.3	25.0	2.92
1986	100.0	144.6	192.5	267.0	266.8	25.0	2.98
1987	100.0	144.1	191.3	264.4	264.0	24.9	2.91
1988	100.0	140.6	183.8	252.2	258.3	25.8	2.97
1989	100.0	139.5	181.4	247.6	254.4	26.0	2.93
1990	100.0	135.7	172.7	231.3	240.1	26.5	2.71
1991	100.0	136.8	175.0	234.1	239.3	25.8	2.62
Mean	100.0	137.6	176.1	231.5	226.6	24.3	2.33

[1] Years of work experience.

[2] This is the partial derivative of the wage equation with respect to experience evaluated at the annual mean of estimated male experience or the percentage increase in earnings due to an additional year of experience.

Source: Calculated from table E-3.

the schooling coefficients for women were lower than those for men in each of the education groups and in all years (with the insignificant exceptions of the high school graduates in 1950 and the 9-11 schooling group in 1973), except for the lowest schooling group for which the coefficients for women were larger. In 20 of the 36 possible cases, the differences are not significantly different from zero, but there are at least two coefficients in each year, except 1968, for which the differences are statistically significant. In 1968, only the estimated coefficient for the interaction term for those with 1-4 years of schooling was statistically significant.

The male-female differences in the estimated coefficients for the experience terms are more consistently significant. The first order experience term is always significantly lower for women than men, while the second order term is twice significantly higher and once significantly lower. This confirms the results of tables E-7 and E-8 which show that the labor price-experience profiles for women are flatter than those for men.

In sum, the above findings, which are consistent with those reported by other researchers, indicate that the returns to schooling and experience differ for men and women as mea-

Table E-8. Indexes of the hourly price of labor by years of work experience, peak earnings years, and marginal increase in hourly labor prices for women, 1950-91

Survey year	Years of work experience					Peak[1] earnings	Marginal[2] increase in hourly labor price (percent)
	0	5	10	15	20		
1950	100.0	113.1	123.1	129.1	130.3	18.8	1.06
1960	100.0	113.2	124.3	132.7	137.4	23.5	1.25
1968	100.0	120.2	135.6	143.5	142.6	17.0	1.23
1969	100.0	115.9	128.1	134.9	135.5	17.9	1.13
1970	100.0	114.9	126.0	131.9	131.8	17.5	1.04
1971	100.0	118.7	132.9	140.2	139.4	17.0	1.21
1972	100.0	121.9	138.9	147.8	146.9	17.1	1.44
1973	100.0	121.3	137.7	146.0	144.9	16.9	1.48
1974	100.0	122.6	140.9	151.9	153.6	18.4	1.85
1975	100.0	121.4	138.3	147.6	147.8	17.6	1.65
1976	100.0	124.0	142.5	151.6	149.3	16.5	1.69
1977	100.0	126.3	146.4	156.0	152.6	16.2	1.84
1978	100.0	123.6	142.5	153.2	153.5	17.7	1.90
1979	100.0	120.3	136.1	145.0	145.5	17.7	1.72
1980	100.0	123.4	141.3	150.2	148.2	16.6	1.76
1981	100.0	123.5	141.5	150.5	148.7	16.7	1.74
1982	100.0	122.1	139.3	148.4	147.6	17.1	1.72
1983	100.0	131.7	156.0	166.2	159.3	15.5	1.92
1984	100.0	127.6	150.0	162.4	162.1	17.4	2.06
1985	100.0	131.9	157.9	171.5	169.1	16.8	2.23
1986	100.0	131.5	157.3	171.2	169.4	17.0	2.25
1987	100.0	132.2	158.6	173.0	171.3	17.0	2.29
1988	100.0	134.8	163.5	178.7	175.8	16.7	2.38
1989	100.0	128.8	152.4	165.5	165.1	17.3	2.06
1990	100.0	131.7	157.5	170.8	168.2	16.7	2.06
1991	100.0	132.8	159.4	172.9	169.5	16.5	2.05
Mean	100.0	124.2	143.4	153.6	152.5	17.3	1.73

1 Years of work experience.

2 This is the partial derivative of the wage equation with respect to experience evaluated at the annual mean of estimated female experience or the percentage increase in earnings due to an additional year of experience.

Source: Calculated from table E-4.

sured by the estimated coefficients. Therefore, it would be inappropriate to estimate a single earnings function from pooled data for the two groups.[27]

Tests for intertemporal changes. The second test is concerned with the intertemporal stability of the earnings functions. Casual observation of the estimated coefficients for schooling and experience and the price of labor reported in tables E-5 through E-8 showed that it is difficult to discern whether the annual movements are merely random fluctuations or real short-term and long-term movements. This next test determines whether the observed changes over time are the result of sampling errors. If the coefficients prove to be

27 This conclusion is further confirmed by joint tests of the coefficients of the interaction terms shown in table E-9. In one test, the six women-schooling interaction terms were omitted; in a second test, the two women-experience interaction terms were omitted. The tabulation below for selected years shows that the "F" statistics were significant in all cases. (The asterisk indicates statistical significance at the 95-percent confidence level.) This means that omitting the subset of coefficients for schooling or experience in any of these years would significantly reduce the explained variation (R^2) in the dependent variable, the natural log of hourly earnings. For a discussion of joint tests of regression coefficients, see Robert S. Pindyck and Daniel L. Rubinfeld (1981), pp. 117-20.

These separate tests are stronger than omitting the schooling and experience coefficients at the same time. If some coefficients are not significantly different, stacked regressions can be used. A joint test would not reveal that only some of the parameters are significantly different.

	1950	1960	1968	1973	1978	1983
Omitting all schooling coefficients	22.9*	4.4*	1.0	2.8*	5.1*	8.0*
Omitting all experience coefficients	381.9*	140.6*	88.7*	145.4*	300.8*	331.7*

Table E-9. Estimated coefficients for schooling, experience, and interaction terms with women, pooled samples of men and women for selected years, 1950-83

Variable	1950	1960	1968	1973	1978	1983
Men						
Intercept	-0.319 (29)	0.053 (3.2)	0.330 (16)	0.529 (24)	0.689 (38)	1.031 (45)
S1TO4	-.166 (19)	-.221 (12)	-.203 (7.6)	-.232 (7.3)	-.210 (7.0)	-.241 (6.0)
S9TO11	.104 (17)	.119 (12)	.102 (7.6)	.089 (5.8)	.155 (11)	.124 (6.6)
S12	.169 (27)	.201 (20)	.234 (19)	.210 (15)	.308 (25)	.306 (18)
S13TO15	.259 (29)	.349 (26)	.353 (22)	.315 (19)	.378 (27)	.372 (20)
S16	.432 (36)	.564 (33)	.520 (28)	.500 (26)	.573 (37)	.560 (30)
S17UP	.410 (23)	.573 (24)	.624 (24)	.569 (23)	.696 (37)	.732 (35)
Experience	.0133 (57)	.0146 (39)	.0137 (27)	.0167 (34)	.0195 (51)	.0197 (44)
Experience squared[1]	-.68 (44)	-.73 (28)	-.70 (21)	-.88 (26)	-1.04 (40)	-1.02 (34)
Interaction terms with women						
Intercept	-.069 (4.4)	-.066 (2.8)	-.204 (6.7)	-.120 (3.7)	-.018 (.7)	-.016 (.5)
S1TO4	.141 (6.7)	.103 (2.4)	.146 (2.3)	.166 (2.5)	.165 (2.7)	.020 (.3)
S9TO11	-.026 (2.2)	-.035 (1.8)	-.015 (.6)	.008 (.3)	-.059 (2.5)	-.025 (.8)
S12	.002 (.2)	-.023 (1.3)	-.013 (.6)	-.014 (.6)	-.066 (3.1)	-.065 (2.3)
S13TO15	-.010 (.6)	-.069 (2.8)	-.040 (1.4)	-.013 (.5)	-.020 (.9)	-.006 (.2)
S16	-.159 (7.3)	-.148 (4.0)	-.035 (.9)	-.068 (1.6)	-.086 (3.1)	-.029 (.9)
S17UP	-.110 (3.6)	-.107 (1.8)	-.056 (.9)	-.113 (2.0)	-.039 (1.1)	-.085 (2.3)
Experience	-.0062 (14)	-.0077 (11)	-.0030 (3.3)	-.0055 (6.1)	-.0071 (10)	-.0033 (4.2)
Experience squared[1]	.22 (6.2)	.36 (6.7)	-.08 (1.0)	.07 (.8)	.17 (2.7)	-.29 (4.0)

NOTE: t-statistics in parentheses.

[1] Coefficient is multiplied by 10,000.

stable, the samples could be pooled and the efficiency of the coefficients increased. Otherwise, the earnings weights must be based on annual measures of the price of labor.

The data for 1968 are pooled with those for 1973, 1978, and 1983, and the earnings function is estimated for the pooled sample. The function includes the variables specified in table E-2; a set of interactions between year-specific (1973, 1978, or 1983) dummy variables and education or experience; and year-specific intercepts. Thus, 1968 is used as the base year, and the coefficients for the interaction terms are estimates of the differences between the coefficients for each of the 3 years and those for 1968. The functions are estimated separately for men and women. The results are in tables E-10 and E-11 (year-specific intercepts are not shown).

In the case of men, there were statistically significant changes between the 1968 schooling coefficients and those for 1978 and 1983. Men with 17+ years of schooling received higher payments for their schooling than their 1968 counterparts (table E-10). In addition, men with a high school diploma in 1983 and men with 9-11 years of schooling in 1978 received higher returns for their education. Thus, there was a statistically significant increase in the schooling coefficient for those with 17 or more years of education, from 0.627 in 1968 to 0.695 in 1978 and to 0.735 in 1983. The first order ex-

Table E-10. Estimated coefficients for schooling, experience, and intertemporal interaction terms, men, selected years, 1968-83

Variable	Base year 1968	Interaction coefficients[1] 1973	1978	1983
S1T04	-0.199 (7.4)	-0.035 (.9)	-0.010 (.2)	-0.040 (.8)
S9T011	.102 (7.6)	-.014 (.7)	.052 (2.6)	.022 (1.0)
S12	.231 (19)	-.022 (1.2)	.074 (4.0)	.073 (3.6)
S13T015	.351 (22)	-.038 (1.7)	.025 (1.1)	.020 (.8)
S16	.520 (28)	-.021 (.8)	.051 (2.0)	.042 (1.6)
S17UP	.627 (24)	-.059 (1.7)	.068 (2.1)	.108 (3.3)
Experience	.0136 (27)	.0031 (4.4)	.0059 (8.9)	.0060 (9.0)
Experience squared	-.000070 (21)	-.000018 (4.0)	-.000034 (7.7)	-.000032 (7.1)

NOTE: t-statistics in parentheses.
[1] Estimated coefficients (and t-score) for interaction terms between schooling and experience in 1968 and successive years.

Table E-11. Estimated coefficients for schooling, experience, and intertemporal interaction terms, women, selected years, 1968-83

Variable	Base year 1968	Interaction coefficients[1] 1973	1978	1983
S1T04	-0.067 (1.2)	-0.004 (.0)	0.014 (.2)	-0.160 (1.9)
S9T011	.090 (4.5)	.010 (.3)	.009 (.3)	.011 (.4)
S12	.227 (13)	-.028 (1.1)	.020 (.8)	.017 (.7)
S13T015	.318 (14)	-.011 (.4)	.042 (1.4)	.049 (1.5)
S16	.487 (14)	-.054 (1.2)	.000 (.0)	.043 (1.0)
S17UP	.568 (9.9)	-.114 (1.5)	.088 (1.3)	.075 (1.2)
Experience	.0108 (15)	.0006 (.6)	.0016 (1.6)	.0056 (5.8)
Experience squared	-.000079 (11)	-.000005 (.5)	-.000008 (.8)	-.000053 (5.6)

NOTE: t-statistics in parentheses.
[1] Estimated coefficients (and t-score) for interaction terms between schooling and experience in 1968 and successive years.

perience coefficients were significantly larger, and the second order experience coefficients were significantly lower in all years after 1968.

In the case of women, none of the schooling coefficients were significantly different from the 1968 coefficients. However, taken as a group, the education coefficients in the later years are significantly different from 1968.[28] The first order experience coefficients increase, and the second order terms decrease over the period, but the changes were statistically different only between 1968 and 1983.

In sum, the results indicate that the earnings functions shifted significantly after 1968, but the changes are greater for men than women. For both men and women, the labor price-experience profiles have become steeper over time. Therefore, the coefficients are not stable over time, and the price of labor will be estimated annually.

[28] The statistical significance of the changes in the schooling and experience coefficients between 1968 and 1983 for both men and women are also confirmed by joint tests of the coefficients of the interaction terms shown in tables E-10 and E-11. The tests are similar to those described in footnote 27. Again, earnings functions were estimated. One omitted the 1973, 1978, and 1983 interaction terms with schooling, and the other omitted the interaction terms with experience. The "F" statistics shown in the tabulation below measure the loss of explanatory power from omitting these interaction terms. (The asterisk indicates statistical significance at the 95-percent level.) Year-specific interaction terms for veteran and part-time workers are also included in the estimated equations. This specification places fewer restrictions on the schooling and experience coefficients and results in a stronger test of intertemporal changes.

	Men	Women
Omitting all schooling coefficients	4.0*	2.6*
Omitting all experience coefficients	26.7*	10.2*

Table E-12. Estimated earnings functions for men, selected years, 1950-91

Variable	1950	1960	1968	1978	1983	1991
Constant	-0.324	0.133	0.345	0.684	1.042	1.276
	(28)	(7.5)	(16)	(34)	(42)	(57)
S1TO4	-.166	-.156	-.199	-.208	-.239	-.111
	(19)	(8.3)	(7.7)	(6.7)	(5.7)	(2.9)
S9TO11	.104	.100	.102	.154	.124	.172
	(17)	(10)	(7.9)	(11)	(6.3)	(9.0)
S12	.168	.179	.231	.305	.305	.317
	(28)	(18)	(19)	(24)	(17)	(19)
S13TO15	.257	.324	.351	.376	.371	.485
	(29)	(25)	(23)	(26)	(20)	(27)
S16	.433	.545	.520	.571	.562	.745
	(37)	(33)	(29)	(35)	(28)	(40)
S17UP	.409	.550	.627	.695	.735	.899
	(23)	(29)	(25)	(36)	(34)	(43)
EXPER	.0133	.0147	.0136	.0196	.0196	.0174
	(57)	(40)	(28)	(49)	(43)	(42)
EXPSQ	-.000069	-.000074	-.000070	-.000104	-.000102	-.000084
	(45)	(29)	(21)	(38)	(33)	(30)
PTIME	-	-	-.316	-.187	-.196	-.274
			(19)	(15)	(15)	(21)
VET	.032	.026	.048	.054	.061	.004
	(6.7)	(3.4)	(5.2)	(6.5)	(6.2)	(.5)
NEWENG	-.150	-.090	-.104	-.029	-.052	.061
	(14)	(5.4)	(5.4)	(1.7)	(2.7)	(3.5)
MIDATL	-.083	-.044	-.064	.008	-.043	.004
	(10)	(3.5)	(4.4)	(.6)	(2.9)	(.3)
ENCENT	-.047	.006	-.020	.048	.019	-.013
	(5.9)	(.5)	(1.4)	(3.9)	(1.3)	(1.1)
WNCENT	-.201	-.159	-.105	-.034	-.047	-.100
	(20)	(9.7)	(5.6)	(2.1)	(2.5)	(6.1)
SATL	-.241	-.256	-.190	-.091	-.125	-.097
	(26)	(18)	(12)	(6.7)	(8.3)	(7.8)
ESCENT	-.289	-.291	-.285	-.089	-.075	-.110
	(24)	(16)	(14)	(5.1)	(3.7)	(6.3)
WSCENT	-.251	-.247	-.191	-.073	-.012	-.103
	(24)	(16)	(11)	(4.9)	(.8)	(7.2)
MOUNTAIN	-.125	.086	-.082	-.032	-.030	-.084
	(8.7)	(3.9)	(3.3)	(1.6)	(1.4)	(4.7)
CCITY	-	-	.088	.049	.037	.010
			(8.3)	(5.2)	(3.3)	(1.0)
BCITY	-	-	.177	.140	.134	.144
			(17)	(16)	(13)	(17)
URBAN	.170	.069	-	-	-	-
	(34)	(8.3)				
R^2	.163	.201	.235	.315	.269	.346
Degrees of freedom	69,660	26,838	24,102	26,269	25,488	25,644

NOTE: t-statistics in parentheses.

Table E-13. Estimated earnings functions for women, selected years, 1950-91

Variable	1950	1960	1968	1978	1983	1991
Constant	-0.378 (21)	0.021 (.7)	0.092 (2.7)	0.678 (29)	1.001 (36)	1.210 (43)
S1TO4	-.025 (1.3)	-.120 (3.0)	-.067 (1.1)	-.053 (1.1)	-.228 (3.6)	.006 (.1)
S9TO11	.081 (7.9)	.087 (5.4)	.090 (4.1)	.099 (5.5)	.102 (4.3)	.137 (5.5)
S12	.173 (19)	.181 (12)	.227 (12)	.247 (15)	.244 (12)	.310 (14)
S13TO15	.252 (20)	.285 (14)	.318 (13)	.360 (20)	.367 (16)	.481 (20)
S16	.275 (14)	.420 (13)	.487 (13)	.487 (22)	.529 (22)	.739 (30)
S17UP	.297 (11)	.469 (8.8)	.568 (9.1)	.657 (22)	.643 (22)	.853 (31)
EXPER	.0071 (18)	.0069 (12)	.0108 (13)	.0123 (23)	.0161 (26)	.0167 (28)
EXPSQ	-.000047 (15)	-.000037 (7.7)	-.000079 (10)	-.000087 (16)	-.000130 (21)	-.000126 (21)
PTIME	-	-	-.084 (5.4)	-.145 (16)	-.178 (19)	-.184 (21)
VET	-	-	-	-	-	-.056 (1.3)
NEWENG	-.099 (6.3)	-.061 (2.5)	-.019 (.7)	-.026 (1.4)	-.018 (.9)	.119 (6.7)
MIDATL	-.053 (4.2)	-.022 (1.1)	-.035 (1.6)	.008 (.6)	-.023 (1.5)	.007 (.5)
ENCENT	-.086 (6.8)	-.066 (3.4)	-.036 (1.6)	-.007 (.5)	-019 (1.3)	-.093 (7.1)
WNCENT	-.214 (14)	-.215 (8.6)	-.152 (5.2)	-.101 (5.6)	-.049 (2.6)	-.115 (7.0)
SATL	-.210 (14)	-.238 (11)	-.111 (4.6)	-.074 (5.0)	-.076 (4.9)	-.098 (7.4)
ESCENT	-.324 (17)	-.331 (12)	-.218 (6.7)	-.113 (5.6)	-.133 (6.4)	-.165 (9.0)
WSCENT	-.339 (20)	-.334 (13)	-.184 (6.3)	-.113 (6.7)	-.074 (4.3)	-.147 (9.6)
MOUNTAIN	-.123 (5.0)	-.180 (5.1)	-.139 (3.6)	-.091 (4.1)	-.064 (2.9)	-.085 (4.5)
CCITY	-	-	.162 (10)	.104 (9.9)	.126 (11)	.113 (11)
BCITY	-	-	.167 (10)	.117 (12)	.139 (13)	.167 (19)
URBAN	.150 (19)	.112 (8.2)	-	-	-	-
R^2	.090	.095	.077	.159	.154	.246
Degrees of freedom	26,915	12,611	12,550	17,650	20,236	22,815

NOTE: t-statistics in parentheses.

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Appendix F: Comparison of Labor Composition Measures

Labor composition has been the subject of empirical study for a number of years.[1] Interest in this field has intensified as speculation over the importance of labor composition changes to the post-1973 productivity slowdown has increased. Labor composition studies have been written by Schultz, Griliches, Chinloy, Denison, and Jorgenson, Gollop, and Fraumeni.[2] The studies by the late Edward Denison and by Dale Jorgenson, Frank Gollop, and Barbara Fraumeni (JGF) are of particular interest because these authors have updated and revised their measures over the years and these two studies represent different approaches to measuring labor composition.

This appendix compares the work of Denison and JGF with the BLS measures presented in this study. The first section compares recent measures of labor composition. The next section examines the sources of worker heterogeneity. The third section discusses the measurement of hours of work and the corresponding weights. The final section estimates the quantitative effect of these differences on the measures of labor composition.

The comparison is limited to labor composition growth only. The measurement of the raw labor input and labor's share of the value of output has, in large part, been discussed in Bulletin 2178, *Trends in Multifactor Productivity, 1948-81*. Also, the impact of labor composition on multifactor productivity is not discussed.[3]

Measures of labor composition growth

Table F-1 contains the indexes of labor composition for BLS, and JGF. The BLS index of labor composition measures changes in the skills of workers arising from changes in the education, work experience, and sex composition of the private business sector work force. Denison's index of labor composition for nonresidential business classifies workers by education, age, sex, and labor force groups.[4] The labor force groups include part-time workers, farm workers, full-time nonfarm wage and salary workers, and full-time nonfarm proprietors and unpaid family workers.[5] Denison used these four groups to measure two additional sources of change in the efficiency of an hour of work caused by changes in the length of the workweek and shifts of full-time employment from farming to full-time employment outside of farming. JGF's measure reflects changes in labor composition across education, age, sex, class of worker (proprietor or employee), and occupation for the total economy. JGF also produce labor composition measures by industry which, of course, requires industry as an additional category.

It is readily apparent that the labor composition measures of Denison and JGF grow more rapidly than the BLS measures through 1979. The growth rates are summarized in table F-2. Denison's measure grew 0.61 percent annually during the years 1948-79, and JGF's measure rose at an annual rate of 0.60 percent, while the BLS measure grew 0.23 percent annually. Table F-1 shows the cumulative effect through 1979 of these differing growth rates: The BLS measure of labor composition rose by 7.29 percent, while Denison's and JGF's measures increased by 20.83 and 20.55 percent, respectively.

It has been suggested that the large influx of women and young workers into the work force in the 1970's has reduced the average experience level of workers and slowed the growth of productivity. For all three measures of labor composition, the growth rate is lower during the 1973-79 period. However, except for JGF, the decline in the labor composition growth rate is small, 0.2 percent or less. The measures of JGF indicate that labor composition growth comes to a halt in the 1973-79 period. This contrasts sharply with the 0.6- to 0.7-percent annual increases in the earlier period. As a result, only JGF find a significant role for labor composition in explaining the productivity slowdown. For BLS and Denison, labor composition has a small effect.[6]

Furthermore, both the BLS and Denison measures indicate that the decline was temporary. Both BLS and Denison show

[1] See R. Solow (1957) and J. Schmookler (1952) for early contributions to the concepts and measurements of labor composition.

[2] See T. W. Schultz (1961), Z. Griliches (1970), P. Chinloy (1981), E. Denison (1985), and D. Jorgenson, F. Gollop, and B. Fraumeni (1987).

[3] In the BLS measures, labor's share of the value of output over the postwar period is approximately 70 percent. The corresponding shares for Denison and JGF are 80 percent and 60 percent respectively. The contribution of labor composition to multifactor productivity can be approximated as the product of the growth rate of labor composition and labor's cost share.

[4] The measures cited in this appendix come from Denison's 1985 volume, *Trends in American Economic Growth, 1929-1982*. The description of Denison's methods also are based on this volume, but the most thorough description is found in his 1974 volume, *Accounting for United States Economic Growth, 1929-1969*. Descriptions of methods from the earlier volume are sometimes used. While every effort has been made to describe Denison's current methods, some revisions between the 1974 and 1985 volumes may have been overlooked.

[5] In Denison's definition, farm workers include everyone working in the farm sector: Wage and salary workers, proprietors, and unpaid family workers.

[6] For Denison and BLS respectively, labor composition accounts for 0.1 and 0.2 percent of the productivity slowdown during the period 1973-79. In contrast, JGF explain 0.4 percent of the productivity slowdown by the labor composition measures for the period 1973-79.

Table F-1. Indexes of labor composition, 1948-90
(1948 = 100)

Year	BLS	Denison	JGF
1948	100.00	100.00	100.00
1949	100.21	101.49	100.12
1950	100.98	101.72	101.43
1951	101.25	102.66	102.63
1952	101.53	104.21	104.54
1953	102.29	105.28	105.62
1954	103.03	106.15	105.97
1955	103.23	106.73	106.21
1956	103.40	106.87	106.81
1957	103.85	108.07	108.12
1958	104.25	109.37	108.60
1959	104.38	110.04	109.56
1960	104.91	110.43	112.66
1961	105.50	111.62	111.71
1962	106.50	111.76	113.74
1963	106.73	112.37	114.10
1964	106.78	112.65	115.29
1965	106.68	112.57	115.65
1966	106.66	113.14	117.20
1967	106.83	114.54	118.16
1968	106.58	115.08	118.88
1969	106.99	115.06	118.68
1970	107.46	116.11	120.07
1971	107.15	116.94	120.19
1972	107.20	117.01	119.47
1973	106.97	117.32	119.95
1974	107.68	118.18	120.55
1975	107.73	120.02	121.74
1976	107.50	120.14	122.10
1977	107.50	120.11	121.74
1978	107.64	120.21	122.46
1979	107.29	120.83	120.55
1980	107.63	121.90	-
1981	108.45	122.91	-
1982	109.56	124.60	-
1983	110.03	-	-
1984	110.15	-	-
1985	110.43	-	-
1986	110.98	-	-
1987	111.26	-	-
1988	112.16	-	-
1989	112.64	-	-
1990	113.26	-	-

Sources: See chapter II, table 19, private business, this bulletin. See Edward F. Denison, *Trends in American Economic Growth, 1929-1982*, Brookings Institution (Washington, DC, 1985). The index is the product of columns 4, 5, 6, and 7 in table 3-1. See also D. Jorgenson, F. Gollop, and B. Fraumeni, (JGF), *Productivity and U.S. Economic Growth*, Harvard University Press (Cambridge, MA, 1987). The index is the fifth column of table 8.1.

Table F-2. Average annual growth rates of labor composition, selected periods, 1948-82
(Percent per year)

Study	1948-79	1948-73	1973-79	1979-82	1973-82
BLS	0.23	0.27	0.05	0.70	0.27
Denison	.61	.64	.49	1.03	.67
JGF	.60	.73	.08	-	-

Source: Table F-1.

that labor composition growth after 1979 was sufficiently rapid to overcome the slow growth rates of the 1973-79 period. As a result, the 1973-82 growth rate equals or exceeds the pre-1973 rate.[7]

The widely practiced principle of last hired/first fired and the steady demand for highly educated workers causes marked fluctuations in the employment of young, female, and less educated workers. As a result, the labor composition measures tend to move counter to the business cycle. Labor composition growth is generally most rapid during the early stages of a recession and slowest during the early phases of recovery. The labor composition measures exhibit differing sensitivity to the business cycle. Table F-3 shows the correlation coefficient between labor composition growth and the growth rate of BLS measures of private business sector hours of work and output.

Denison's measure showed a very strong countercyclical pattern with respect to both hours and output. The BLS measure indicates a weaker countercyclical pattern. The correlation coefficient with hours of work is -0.40, but the correlation with output is an insignificant -0.22.[8] The JGF measure does not demonstrate any significant countercyclical pattern.[9]

Table F-3. Correlation coefficient between the growth rate of various labor composition measures and private business hours and output, 1949-79

Study	Correlation coefficient with:	
	Growth rate of hours	Growth rate of output
BLS	-0.40[1]	-0.22
Denison	-.87[1]	-.64[1]
JGF	-.17	.05

[1] Significant at the 95-percent confidence level.

Differences in research design

This section focuses on the theoretical aspects of the measures, specifically, how the authors differ in identifying sources of productivity differences between individuals. The next section describes the methodological choices made in each study to implement the program of labor composition measurement. Each of these studies is quite complex, so it is only possible to identify the more fundamental differences.

All the measures of labor composition begin with the assumption that workers are paid the value of their marginal

[7] In this bulletin, growth rates are usually given for periods spanning business cycle peaks (for example, 1973-79). In this case, a trough year, 1982, is used because it is the last year in Denison's study. A comparison of peak-to-peak (JGF) with peak-to-trough (Denison) growth rates can be misleading because the difference in growth rates of the two measures also includes the effects of variations in the business cycle. However, the BLS growth rate of labor composition is nearly identical for 1973-82 and 1973-90. This suggests that both the Denison and BLS measures of labor composition do not show a permanent slowdown after 1973.

[8] For the periods 1949-82 and 1949-90, the BLS measure has a significant correlation with output of -0.32 and -0.31, respectively.

[9] JGF's measures of labor composition include government and nonprofit institutions. These sectors exhibit very little cyclical variation in the composition of employment; therefore, JGF's measures would be expected to show less correlation with the business cycle.

product, and therefore worker productivity differences can be distinguished on the basis of hourly compensation. All three studies identify a set of worker traits and then collect the hours and earnings weights required to measure labor composition. BLS uses the worker traits of education, work experience, and sex. Denison differentiated hours on the basis of education, age, sex, and labor force groups. Changes in the workweek of labor force groups are used to estimate the effect of changes in the length of the workweek on the efficiency of an hour's work. The change in total hours arising from shifts in full-time employment between labor force groups is measured and included in Denison's measures of labor composition.[10] JGF use five traits: Education, age, sex, employment class, and occupation.[11]

Unlike the other studies, Denison measured labor composition as the product of four separate indexes: Education, age-sex, intergroup shifts, and the efficiency offset or intragroup changes. This implies that the earnings weights can be derived independently. BLS and JGF measure the earnings weights jointly for a combination of traits, and no assumption of independence is required. If Denison was correct, the contribution of education to labor input and productivity, for example, could be measured directly. JGF create partial indexes which may indicate the contribution of a single characteristic to labor input. However, these partial indexes are derived under the assumption that labor is homogeneous with respect to all but one characteristic even though JGF assume labor is heterogeneous across all five traits. BLS estimates of partial indexes can be found in appendix H. These indexes are derived by decomposing the wage model of appendix E so that the returns to only one trait at a time are considered in each partial index. This treatment of the wage model permits the partial indexes to generally maintain the assumption of homogeneity as required. Even with these assumptions, the derivation of these partial indexes does not strictly satisfy an exact decomposition of labor composition into the contribution of each trait.[12]

Occupation and industry. Denison explicitly chose not to differentiate workers by their occupation. "Education affects earnings both by improving an individual's ability to perform in a particular type of work and by enlarging the range of occupations in which he is able to perform at all effectively. Both are properly counted as results of education that we do not wish to remove. There is no reason to standardize the earnings data by industry or occupation...."[13] He further stated, "An improvement *over time* in the allocation of workers (classified by age and sex or by other characteristics) among jobs could raise the average productivity of all workers and hence output. If such a change occurred I would not wish to classify the gain as a contribution of labor input but instead as a contribution of output per unit input (specifically, of improved resource allocation)."[14]

BLS does not include industry or occupation in its measures. As stated in appendix A, the focus of this study has been narrowed to gain an understanding of the sources of labor composition change. Inclusion of occupation and industry would make interpretation of the sources of labor composition change more ambiguous because shifts in occupational or industrial wage differentials may also reflect shifts in the education or experience of the work force. Such relative pay shifts lead to corresponding labor composition changes. JGF include only occupation in their labor composition measure. The authors consider labor composition growth arising from industrial wage differentials to be more properly termed a gain in resource allocation.[15]

Class of worker and labor force groups. BLS does not distinguish hours by class of worker. Earnings weights are based on the price of labor for employees only, and proprietors and unpaid family workers are assigned a price of labor equal to an employee of the same characteristics. Earnings of proprietors include a return to both labor and capital, and debatable assumptions are required to disentangle the two sources of income. JGF assume the return to capital is identical in all sectors. Labor earnings of proprietors are then the residual income after subtracting payments to capital. Unlike BLS which assigns equal earnings weights to all employment classes, shifts in hours between employees and proprietors cause changes in JGF's measures of labor composition. Denison assumed the equality of the annual labor inputs of three full-time groups, farm workers, nonfarm proprietors and unpaid family workers, and nonfarm employees.[16] However, the shift toward nonfarm employment tends to cause a decline in measured hours of input because nonfarm employees work fewer hours per year. Since by Denison's assumption, labor input is unchanged by these shifts, the final index of labor input should also be unaffected. Denison achieved this result by an index of intergroup shifts. In this way, he offset the decline in total hours from shifts to nonfarm employment by an increase in the intergroup index and ultimately the index of labor composition.

Education. In principle, JGF and BLS treat education in the same manner. However, Denison adjusted his education weights for ability and socioeconomic status. The adjustment is small except for persons with 5 years or more of college. Nei-

[10] Denison calls the changes in the efficiency of an hour of work the Index of Intragroup Changes, and the employment shift between groups is labeled the Index of Intergroup Shifts.

[11] JGF also use a sixth trait, industry, when measuring labor composition at the industry level.

[12] See appendix H for a discussion of the limitations on the interpretation of partial indexes. For a more complete discussion, see Rosenblum, Dean, Jablonski, and Kunze (1990).

[13] Denison (1974), *Accounting for United States Economic Growth, 1929-1969*, p. 227.

[14] Ibid, p. 33. (Italics in original.)

[15] Among possible interpretations, industrial wage differentials have been frequently interpreted as either a misallocation of resources or a reflection of unmeasured labor quality differences. JGF implicitly ascribe at least part of the differential to a misallocation of resources (p. 314).

[16] Denison (1974), *Accounting for United States Economic Growth 1929-1969*, pp. 30, 42.

ther BLS nor JGF make similar adjustments. Furthermore, Denison used male educational earnings weights for all persons, while BLS and JGF use separate weights for men and women.

Work experience and age. BLS classifies hours by the level of work experience rather than age, as is done in the other studies. Ultimately, all studies are attempting to measure the level of on-the-job training. BLS prefers work experience because it is more closely related than age to the concept of training. In appendix C, this was confirmed by comparing the coefficients of the estimated wage models based on age versus years of actual experience. In wage models using age as an explanatory variable, the effect of education on earnings is overstated, and the effect of age or potential experience is understated. For this reason, BLS developed an equation to measure work experience based on age, education, and other traits. As a result both the hours of work and the earnings weights are measured in terms of work experience instead of age.

Part-time workers and intragroup changes. BLS and JGF use equal earnings weights for all workers regardless of their work schedule. Denison did not. Instead, he made two assumptions. First, he assumed that the lower hourly wage for part-time workers reflects lower productivity. He assumed that the hourly earnings of part-time workers are only 80-90 percent of full-time workers and adjusted the earnings weights of part-time workers in his age-sex composition index accordingly.

Second, Denison assumed that persons working many hours per week suffer a decline in their productivity. As a result, a decline in the workweek for such persons would lead to a less than proportionate decline in their labor input. That is, the efficiency of an hour of work rises as the length of the workweek falls. For full-time nonfarm employees, a reduction of 1 hour in the average work week results in an offset of 30 percent. In effect, only 0.7 hours of labor input are lost. He further assumed a 100-percent offset for farm workers and for nonfarm proprietors and unpaid family workers. This implies that labor input for these workers is independent of the length of the workweek. For part-time workers, no offset is assumed which implies that the efficiency of an hour is unaffected by the length of their workweek. Based on these assumptions and changes in the workweek, Denison developed an index of "intragroup changes" that incorporates all these efficiency offsets.

Race and sex. Finally, all three studies classify hours by sex, and none of them classifies hours or earnings by race. No adjustment to the earnings weight of men or women is made by BLS or JGF, but Denison used only the earnings of men to measure the impact on labor input of changes in educational attainment of men and women.

Differences in methods

Hours. The construction of the hours matrices begins with an enumeration of the possible types of labor. As can be seen in table F-4, each study uses a large number of worker types. BLS identified 1,008, Denison used 576, and JGF classified hours into 1,600 categories.[17]

Table F-4. Classification system of hours worked by BLS, Denison, and Jorgenson, Gollop, and Fraumeni

Trait	Number of categories for each trait			
	BLS		Denison	JGF
	Initial	Final		
Total	4,536[3]	1,008	576	1,600
Education	7	7	9	5
Age	72	-	8	8
Work experience	-	72	-	-
Sex	2	2	2	2
Class of worker[1]	-	-	-	2
Labor force group[2]	-	-	4	-
Occupation	-	-	-	10
Marital status	2	-	-	-
Children	4	-	-	-

[1] Includes proprietors and employees.

[2] Includes part-time workers, all farm workers, full-time nonfarm wage and salary workers, and full-time nonfarm proprietors and unpaid family workers. Denison used these groups when computing indexes of intragroup changes and intergroup shifts.

[3] Men are not classified by marital status or number of children. There are 504 categories for men and 4,032 for women.

The decennial census and the Current Population Survey (CPS) are the principal sources of data for all three studies. The census and the CPS are the only sources of hours and earnings data cross-classified by worker characteristics for the entire postwar period. All three studies prefer the establishment-based survey for their measures of hours, but because this survey lacks the required detail of employment and earnings by trait, the labor composition measures use the household surveys. Since all three studies reconcile their hours measures to the establishment survey but estimate the distribution of hours using the household survey, all studies maintain the assumption that the distribution of hours by trait is the same in both the household and establishment surveys.

BLS measures labor composition for the private business sector. The hours of workers in government, nonprofit institutions, and private households are excluded. Denison measured labor input for the nonresidential business sector which also excludes the hours of government, nonprofit institutions, and private households. Jorgenson, Gollop, and Fraumeni measure labor composition for the total economy though industry measures are also available.

[17] JGF also measure labor composition for 51 industries. These more detailed measures of labor composition use a total of 81,600 categories of labor.

The BLS private business sector hours matrices for workers age 14 or older are taken from computer-coded individual responses to the 1950 and 1960 decennial censuses and the March CPS of 1968-91. From these data, the product of employment, average weekly hours, and weeks worked per year are initially cross-classified by education and age for men and by education, age, marital status, and the number of children born for women. For the remaining years, matrices of hours are initialized based on the distributions of hours in the decennial census and the CPS and are generated from less detailed marginal distributions using the RAS multiproportional interpolation method as described in appendix D. BLS uses an experience equation, described in appendix C, to condense this classification system into the 1,008 worker categories cross-classified by education, work experience, and sex. The construction of the hours matrices is described in detail in appendix D.

JGF develop measures of hours for each industry, and this requires hours matrices based on jobs rather than persons. JGF create a matrix of jobs controlled to the BEA establishment estimates. Average weekly hours per job are also reconciled to establishment-based measures of hours in the national accounts. The total hours matrix is the product of these two matrices and covers workers age 14 or older in the total economy.

JGF did not have access to the individual CPS or census responses, and as a result they could not avail themselves of cross-classified data. Instead, both the employment and average weekly hours matrices must be estimated from published census and CPS marginal distributions using the RAS multiproportional interpolation method.[18] The estimated matrices for the decennial census years are created first, and these matrices are then used to initialize matrices in the remaining years. While the decennial census samples are sufficiently large to create reliable matrices, the marginal distributions in other years use CPS samples of approximately 50,000 households to measure the hours of 81,600 categories of workers at the industry level.[19] At the level of the total economy, only 1,600 kinds of workers are required, and the CPS should be large enough to reliably estimate the hours in each cell.

Denison measured labor composition based on a labor input series of persons rather than jobs. Like JGF, Denison did not use computer-coded survey responses and worked from published and unpublished tabulations. Denison calculated four separate indexes, education, age-sex, intergroup changes, and the efficiency offset, to measure labor composition rather than cross-classify workers by all traits simultaneously. Denison's main data sources did not always provide enough detail, and some estimation was required, based on ancillary data sources. Furthermore, these data had to be converted to full-time equivalents to make use of the available earnings data.

Earnings weights. All studies weight the growth in hours based on the relative earnings of each group of workers. Only the earnings of workers included in the measures of hours are used. The BLS earnings weights for the private business sector exclude government and private household workers, but employees of nonprofit institutions are not explicitly identified in the data. Only employees in nonprofit welfare organization are excluded; their occupation identifies them. As with the BLS measures, Denison's earnings weights also exclude workers in government, private households, and, when possible, nonprofit institutions. JGF include all workers in their earnings weights for the total economy.

The BLS hourly earnings weights are based on wage models estimated annually using survey data from 1950, 1960, and 1968-91. Weights for other years are derived from linear interpolation of the estimated parameters. The wage models are based on samples of individual responses to the household surveys and the census. Only data for paid employees in the private business sector are used. The earnings measure is the ratio of annual wage and salary income before taxes to annual hours of work. This hourly wage measure excludes supplements and payments in kind. As indicated in appendix C, wage models which use age or years since leaving school result in biased estimates of parameters. The wage models in this study use work experience, and the coefficients indicate that the return to schooling is lower and the return to work experience higher than is usually found in other studies.

The need for consistency with the hours matrices requires JGF to measure hourly earnings per job. The matrix of annual compensation per job is the product of annual compensation per person and the number of jobs per person. Annual compensation per job is divided by the previously estimated matrix of annual hours per job. The resultant matrix measures hourly compensation per job. None of the required 81,600 (1,600 for the aggregate labor composition measure) cell means are directly available to JGF. The RAS multiproportional interpolation method generates the matrices for annual compensation per person and for the number of jobs per person.

The JGF compensation measures differ from BLS and Denison in two important ways. First, JGF measure compensation in terms of jobs rather than persons. Second, JGF use compensation including supplements. The household surveys do not include supplements to cash earnings. JGF directly adjust the annual compensation matrix to reflect the statutory employer contributions to Social Security and the unemployment insurance funds. The marginal distributions are controlled to the annual compensation measures from the National Income and Product Account by industry which includes all supplements. JGF note that employer contributions to Social Security and the unemployment insurance funds comprise 70 percent of all supplements. These supplements

[18] JGF state that the marginal distributions of employment are two-, three-, or four-way classifications by traits. For example, employment by age, sex, and education is usually available, but the sample is only for persons age 16 or older.

[19] The estimation of more cells than observations is equivalent to the assumption that many of the parameters (or interaction terms) implicit within the RAS method are zero.

are approximately the same fixed percentage of wages for all employees, but a different percentage for the self-employed. The relative distribution of hourly earnings before and after adjustment for supplements is not likely to differ greatly.[20]

Denison derived four sets of earnings weights, one for each index. The validity of using separate indexes rests on the assumption that the price of an individual trait or skill is independent of the price or quantities of other traits. Neither BLS nor JGF make this assumption. At a minimum, derivation of separate indexes requires detailed cross-classification of earnings consistent with the assumptions of the labor composition study. That is, the earnings weights for one trait must at least control for the other traits presupposed to affect worker productivity. In Denison's case, the traits used to measure intragroup and intergroup shifts can be ignored, and worker productivity depends on age, education, and sex.[21] In developing his education weights, Denison used a six-way classification of earnings by education, age, region, sex, race, and employment class. The development of Denison's earnings weights by education met the criterion of controlling for at least the age and sex of the worker. However, the age-sex composition index was developed using earnings which ignore differences in the educational attainment of workers. Since older workers have less education than younger workers, the earnings of older workers not only reflect their experience but also their lower level of schooling.

There are a number of other factors which influenced Denison's earnings weights. First, published tabulations of earnings include proprietors. Since Denison could not exclude proprietors, the earnings weights implicitly assume that all of a proprietor's income could be attributed to labor. Since proprietors' incomes increase faster with both age and education than wage and salary income, Denison gave greater weight to older workers and more educated workers than he would have done if he had excluded proprietors' earnings from his weights.[22]

Denison began with annual earnings exclusive of supplements. These are converted to full-time and full-year equivalent cash earnings to control for worker differences in the length of the work year. This procedure is accurate provided that full-time workers (those working at least 35 hours per week) have the same average weekly hours. Denison determined that the average workweek of full-time, year-round workers did not vary much by age, and he did use separate measures of average weekly hours for men and women. However, he cannot adjust his earnings weights to reflect the shorter average workweek of less educated full-time and full-year workers.[23] As a result, full-time equivalent earnings yield greater earnings differentials by education level than do hourly earnings, and the weights on the hours of more educated workers are relatively higher than they would be if hourly earnings were used.[24]

Finally, Denison reduced his education earnings weights to control for the effects of ability and socioeconomic status. The adjustment was generally small. With the only important exception being workers with 5 or more years of college, Denison estimated that the ability and socioeconomic status account for less than 11 percent of the marginal increase in earnings from education.[25] Denison estimates that for persons with 5 or more years of college who earn 30 percent more than college graduates, ability accounts for more than a third of the marginal increase in earnings. After controlling for ability and socioeconomic status, Denison only modestly increased the earnings weight of less educated workers and modestly decreased the weight given more educated workers.[26]

Once hours and earnings weights have been developed, BLS and JGF use Tornqvist aggregation (with moving earn-

[20] Empirical tests of this issue can be found in appendix B. The findings of other researchers appear in appendix E. All these studies support the conclusion that compensation share weights would be empirically equivalent to wage share weights in the Tornqvist index number formula.

[21] In Denison's framework, labor force groups can be ignored because the intergroup and intragroup indexes measure changes in labor input due to a reallocation of resources, and these indexes thus use employment not earnings weights.

[22] For full-time year-round workers, the March 1970 CPS indicates that male and female employees with at least a college degree earn 49.7 and 60.6 percent more than their respective counterparts with 9-11 years of schooling. The corresponding figures for self-employed men and women are 155.3 and 104.1 percent. Peak earnings for male and female employees are 61.8 and 20.4 percent above the earnings of their respective counterparts age 20-24. For the self-employed, the figures are 158.5 and 200.0 percent.

[23] For men, the average weekly hours of full-time, year-round workers increase with schooling especially at the highest schooling group. For women, average weekly hours do not vary with schooling except at the highest schooling level. The March 1977 and 1986 CPS average weekly hours of persons working at least 50 weeks per year and 35 hours per week in the nonfarm private business sector are:

	1977		1986	
Schooling	Men	Women	Men	Women
0-4 years	43.0	40.8	42.3	41.8
5-8 years	43.6	40.7	43.3	41.2
9-11 years	44.1	40.8	43.2	41.2
12 years	44.2	40.6	43.8	40.9
13-15 years	44.9	40.6	44.5	41.7
16 years	45.4	40.8	45.7	42.5
17 or more	48.4	43.8	47.7	44.0

[24] JGF follow a similar procedure because their hourly compensation matrix is generated without a marginal distribution which indicates how average weekly hours vary by educational attainment. Furthermore, there is no marginal distribution for the RAS method to indicate how the number of jobs per person varies with education.

[25] The other two groups are persons with 5-7 and 8 years of education. For these two groups, ability accounts for about 20-25 percent of the marginal increase in earnings. However, the total increase in earnings for these two groups is so small that the absolute contribution of ability to earnings is just 1-2 percent.

[26] Denison's 1959 weights before and after adjustment are:

Years of schooling	Before	After
None	71.6	75
1-4	86.5	89
5-7	95.5	97
8	100.0	100
9-11	112.3	111
12	127.3	124
13-15	153.9	147
16	201.3	189
17 or more	264.2	219

Source: Table I-13, *Accounting for U.S. Economic Growth, 1929-1969*.

ings weights) to measure labor composition. Tornqvist aggregation uses compensation share weights which change annually. Denison used three sets of weights for the more than 30 years of his study. Denison did not disagree in principle with moving earnings weights. He believes, however, that the data from the CPS are based on too small a sample to meaningfully measure changes in earnings weights. Instead, the education index weights for 1948-69 are derived from the decennial census of 1960, and the 1970 decennial census is the source of education weights in the remaining years. Similarly, the age-sex composition weights for the period 1948-69 are based on the 1967 CPS earnings distributions. The 1970-75 weights are based on an average of the 1970-75 CPS distributions, and the remaining weights are based on an average of the 1975-79 CPS age-sex earnings distributions.[27]

Quantitative effects of differing methods

This section provides a rough estimate of how theoretical and methodological differences affect the measures of labor composition. The differences are complex, and it is rarely possible to isolate and quantify a single difference in methods. Use caution with regard to the figures cited in this section; most of them are estimates.

Both Denison and JGF include additional traits beyond age, education, and sex. JGF find employment class and occupational changes both contribute positively to labor composition growth. Similarly, Denison found that intergroup changes and the efficiency offset due to changes in the workweek contribute to labor composition growth. Once these traits which are not included in the BLS study were removed, the differences in the labor composition growth rates were greatly reduced. (Appendixes A and E explain why BLS excludes these traits from its measures.)

For Denison, exact estimates of the effects of using the additional traits were possible, but not for JGF. We define the contribution of these traits in the JGF study as the difference between the actual labor composition annual growth rate and the growth rate which would have occurred if the appropriate single trait had been omitted.[28] That is, it is possible to determine what the growth rate of labor composition would have been had JGF not included a specific trait.[29] According to table F-5, these additional traits for Denison and JGF account for about a third of total labor composition growth. However, even after eliminating these additional sources of growth, the BLS measure grows 0.1-0.2 percent less annually.

The remaining differences between BLS and JGF can be attributed to differences in the sectors of the economy studied, the measurement of hours, and the age-sex-education earnings weights. JGF include nonprofit institutions, private household workers, and Federal, State and local government workers. Exclusion of these sectors, as in the BLS study, would raise the JGF growth rate of labor composition because JGF find that labor composition in these sectors grows less rapidly than average. If JGF had adopted the BLS procedure of using only employees in the measurement of labor composition, a slower growth rate would result. (An estimate of the quantitative effect of this assumption for Denison appears below.) In addition, JGF's measure of labor composition growth would have been lower had they been able to incorporate differences in the work year by educational attainment into their hourly compensation matrix. (Again, see below.) Finally, it has been indicated repeatedly that the BLS wage models give greater weight to work experience and less to education than other wage models. Since the average level of experience has fallen and the average level of educational attainment has risen over the postwar period, it is not surprising that the BLS labor composition measure grows less than those of the other studies.

Table F-5. Sources of BLS, Denison, and Jorgenson, Gollop, and Fraumeni labor composition growth, 1948-79 and 1948-82

Item	Annual percentage growth rate			
	1948-82 Denison	1948-82 BLS	1948-79 JGF	1948-79 BLS
Labor composition	0.65	0.27	0.60	0.23
Less contribution to labor composition of:				
Occupation	-	-	.04	-
Employment class	-	-	.08	-
Labor force groups				
Intergroup shifts	.09	-	-	-
Efficiency offset	.11	-	-	-
Interaction effects	-	-	.09	-
Equals:				
Contribution of age, education, and sex to labor composition	.45	.27[1]	.39	.23[1]

[1] BLS measures the labor composition index for experience, education, and sex.

Sources: See table 17, chapter II, this bulletin. See Denison, table 3-1, *Trends in American Economic Growth, 1929-1982*. See also Jorgenson, Gollop, and Fraumeni, table 8.4, *Productivity and U.S. Economic Growth*.

Further comparison of Denison's age-education-sex index with the BLS measures of labor composition is possible. Because the BLS wage model is a semilog form, it is simple to substitute other sets of education and age-sex earnings

[27] See Denison (1985), *Trends in American Economic Growth, 1929-1982*, pp. 64-65.

[28] As another method for measuring the contribution of a single trait, JGF define the main effect of a trait as its first order partial index. This is measured by aggregating the matrices of hours and compensation over all but the trait of interest and recalculating the Tornqvist index. For the 1948-79 period, the annual growth rates for the first order partial indexes of occupation and employment class are 0.30 percent and 0.09 percent, respectively. The interaction effect of the two traits grows 0.00 percent annually. Subtracting these partial indexes from the labor composition growth rate yields an estimate of the contribution of age, sex, and education of 0.21 percent. This estimate is barely half the size of the estimate of 0.39 percent in table F-5.

[29] Note that if JGF had set out to ignore these traits at the beginning of their study, the resultant matrices of hours and earnings would not have been the matrices actually used to generate the partial indexes. This is strictly the result of the RAS interpolation rather than any intentional choice of the authors. Therefore, it is not possible to determine *exactly* the effect of omitting an individual trait, but the marginal contribution should be a good approximation.

weights which more closely correspond to Denison's earnings weights. Beginning with Denison's own age-sex-education index, it is possible to roughly estimate the effects of some of the differences between Denison's and BLS' methods. Table F-6 starts with Denison's own estimates.

Table F-6. Reconciliation of BLS and Denison labor composition growth rates, 1948-82

	Annual growth rate (in percent)	
1) Denison's age-sex-education labor composition indexes[1]	0.45	
Fixed weight labor composition measures		
2) Denison's male education weights, BLS hours and experience-sex weights	.38	
Difference due to hours, experience-sex weights, and index number formula ((1) - (2))		.07
3) BLS male education weights, BLS hours and experience-sex weights	.29	
Difference due to education weights ((2) - (3))		.09
4) BLS male/female education weights, BLS hours and experience-sex weights	.26	
Difference due to separate male/female education weights ((3) - (4))		.03
5) Tornqvist index of labor composition[2]	.27	
Difference due to shift from fixed to Tornqvist Weights ((4) - (5))		-.01

[1] Denison's measure. See table F-5.
[2] BLS measure. See table F-5.

The next entry (line 2) is an attempt to replicate Denison's measure using his education weights with the BLS matrix of hours and experience-sex weights. The result is a hybrid measure of labor composition which grew 0.38 percent annually.[30] The difference of 0.07 percent, shown in table F-6, between Denison's measure and this hybrid measure can be attributed to differences between the two studies in the measures of hours, the use of the predicted experience model, the experience-sex weights, and the choice of index number formulas.[31] Next, by substituting BLS male education weights (line 3) for Denison's (male) education weights for the same 2 years used by Denison, the effect of switching between Denison's weights and the BLS education weights can be estimated. This difference of 0.09 percent in the annual average growth rate is again shown in table F-6.[32] BLS uses both male and female education weights unlike Denison who uses only male weights. The effect of switching from male to male and female education weights (line 4) reduces labor composition growth only 0.03 percent. Finally, BLS uses annually changing wage rates to calculate compensation shares, while Denison used two sets of education weights and three sets of age-sex weights. A switch from fixed to moving weights (line 5) completes the analysis of these differences and yields the BLS measure of labor composition. Note that the three sets of age-sex and two education weights are sufficient to approximate moving annual weights.

As discussed above, use of the BLS education weights (line 3) for men yields a slower rate of labor composition growth than the weights used by Denison (line 2). Three factors contribute to this effect: The exclusion by BLS of self-employed workers, the BLS use of hourly wage rates rather than full-time equivalent earnings in computing the education weights, and the BLS substitution of predicted work experience for age in the wage model. All three factors contribute to smaller earnings differentials between highly educated and less educated workers in the BLS education weights.

It is possible to roughly approximate how the first two factors affect labor composition.[33] Self-employed workers can be included in the sample, and new male education weights derived from a re-estimated wage model. Similarly, the current BLS male education weights can be modified to ignore differences in the average workweek of full-time, year-round workers by educational attainment. In essence, a single standard workweek substitutes for the actual workweek of all full-time, year-round workers. This substitution yields male

[30] The hybrid measure uses Denison's education weights in place of the BLS education parameters of the wage model. Denison's 1959 weights are used through 1969, and the 1969 weights are used for the remainder of the sample.

We also follow Denison's use of fixed age-sex weights for consistency. The experience-sex weights for the period 1948-69 are fixed at the 1967 BLS parameters. Denison uses an average of several years of CPS data for his earnings weights in the remainder of the period. The present comparison uses the 1973 BLS experience-sex parameters for the period 1970-75 and the 1979 BLS parameters for the remainder of the period.

[31] The substitution of estimated experience for age has two subtle effects on labor composition. First, the use of estimated experience leads to a different classification scheme of hours worked. Even assuming away all other differences between the BLS and Denison studies, the growth rate and compensation shares for each type of labor input will differ. This occurs because the distribution of hours by estimated experience is not the same as the distribution by age. Next, both the education and experience weights are affected by the choice of a model for experience in the wage model. Consequently, the effect of the estimated experience model on the labor composition measures is partly included in the difference between Denison's age-sex-education index (table F-6, line 1) and the hybrid fixed weight index which combines Denison's education weights and the BLS hours and experience-sex weights (line 2). However, in table F-6, the difference (table F-6, line 2 - line 3) of 0.09 percent due to education weights also includes the effect on the BLS education weights of using estimated experience in the wage model.

[32] Since Denison's education weights are based solely on the earnings of men, only male BLS education weights are used in the hybrid measures.

[33] To estimate the effect of the estimated work experience model on the labor composition measures would require that the wage models be re-estimated using age as the measure of experience. Furthermore, the hours matrix would have to be reclassified to reflect the age distribution rather than estimated experience. These changes would require a computational burden almost as great as the present study.

education weights based on full-time equivalent (FTE) weekly earnings rather than the BLS procedure of using hourly earnings.[34] Finally, these two effects can be combined by modifying the education weights of all workers including the self-employed by substituting FTE weekly earnings for hourly earnings.[35]

Table F-7 shows how these modifications to the BLS male education weights affect labor composition growth. The effect of switching from hourly wage rates to FTE weekly earnings (line 3 - line 2) increases labor composition growth 0.06 percent annually. The inclusion of self-employed workers (line 4 - line 2) also raises labor composition growth 0.08 percent annually. The combined effect for the two factors (line 5 - line 2) of 0.15 percent annually approximates the sum of the individual effects. Furthermore, the combined effect more than accounts for the difference between the BLS (line 2) and Denison (line 1) labor composition growth rates using fixed male education weights.

The differences shown in table F-7 are likely to overstate the effect of the two factors. The BLS education weights using hourly wages permit the workweek of full-time, year-round workers to increase with education as observed. The BLS male education weights based on FTE weekly earnings use a single workweek for all workers (not just full-time, year-round workers) regardless of educational attainment. However, the adjustment to derive the education weights for FTE earnings is based on full-time, year-round workers only.

Unless the relationship between education and the average workweek is the same for all workers including part-time and part-year workers, an adjustment to the education weights based on the average workweek of only full-time, year-round workers is inexact. Including part-time or part-year workers would likely narrow the difference between the hourly and FTE weekly education weights. As a result, the BLS education weights based on FTE weekly earnings probably uses too large an adjustment (compared to Denison). In addition, the re-estimated wage model for all workers includes only self-employed persons with positive self-employment income. Denison used group averages rather than wage models to derive his weights and includes all self-employed workers regardless of income. Adding persons with self-employment losses to the sample would increase the measure of labor composition in line 4 if the likelihood of losses decreased with education. However, the actual effect of excluding proprietors with losses on the education weights in line 4 is unknown. Nevertheless, it appears based on table F-7 that the

Table F-7. Further reconciliation of BLS and Denison fixed male weight labor composition growth rates, 1948-82[1]

	Annual growth rate (in percent)	
1) Denison's male education weights[2], BLS hours and experience-sex weights	0.38	
2) BLS male education weights[2], based on hourly earnings of employees	.29	
Difference between Denison and BLS[2] education weights ((1) - (2))		0.09
Individual effects of different methods on labor composition		
3) BLS male education weights based on FTE weekly earnings of employees	.35	
Estimated difference from BLS weights due to switch from hourly to FTE earnings ((3) - (2))		.06
4) BLS male education weights based on hourly earnings of employees and self-employed	.37	
Estimated difference from BLS weights due to inclusion of self-employed workers ((4) - (2))		.08
Combined effect of different methods on labor composition		
5) BLS male education weights based on FTE weekly earnings of employees and self-employed	.44	
Estimated difference from BLS weights due to switch to FTE weekly earnings and inclusion of self-employed workers ((5) - (2))		.15

[1] All index growth rates use BLS hours and experience parameters.
[2] Source: Table F-6.

[34] It is possible to re-estimate wage models using weekly rather than hourly wage rates. This is beyond the scope of this study. Instead, an index of average weekly hours for male, year-round, full-time employees (at least 50 weeks per year and 35 hours per week) in the nonfarm private business sector by educational attainment is used to adjust the education parameters.

Schooling group	March 1970 average weekly hours	Index of hours
0-4 years	44.59	98.02
5-8 years	45.49	100.00
9-11 years	45.92	100.95
12 years	46.41	102.02
13-15 years	46.65	102.55
16 years	47.62	104.68
17 or more years	51.94	114.18

The weekly hours of persons with 5-8 years of schooling replace the actual weekly hours of all persons. This substitution alters hourly earnings. A worker's compensation remains the same, but his or her hours have changed. Hourly earnings change in an equal but opposite percentage to the change in weekly hours. The index of hours, therefore, measures the change in earnings resulting from the substitution of actual weekly hours for the workweek of persons with 5-8 years of schooling. The product of the index of hours and the unadjusted education weights yields education weights adjusted to reflect a single workweek for all persons. The adjusted 1970 male education weights are shown below.

Schooling group	Unadjusted weight	Adjusted weight
0-4 years	81.45	79.84
5-8 years	100.00	100.00
9-11 years	114.65	115.74
12 years	127.57	130.15
13-15 years	141.82	145.44
16 years	173.96	182.10
17 or more years	190.61	217.64

The 1960 education parameters are adjusted by the same 1970 index since the quality of 1960 average weekly hours data is much poorer. Finally, the logarithm of the adjusted weights in the semilog wage model is the appropriate adjustment.

[35] The education weights from the wage model which includes the self-employed are multiplied by the same index of hours found in footnote 34. This is strictly correct only if variations in the workweek of full-time workers are independent of the likelihood of being self-employed.

treatment of self-employed workers and FTE weekly earnings accounts for most if not all of the difference between the fixed male education weighted labor composition measures of Denison and BLS.

The labor composition measures of JGF share some similarities with Denison's measures. First, JGF use separate earnings for employees and proprietors. Using the estimate in table F-5 as a guide, JGF's measure of labor composition growth would be 0.08 percent lower if a single earnings weight were used for all workers. However, this measures the impact of switching from two weights to the average weight. BLS omits proprietors and unpaid family workers from the wage model. Instead, these workers are assigned the wage rate of employees with identical traits. The best comparison of the JGF and BLS composition measures would include not only the measured impact in table F-5, but also the effect of switching from an average earnings weight for all persons to an earnings weight of employees only. Comparing Denison's and BLS' measures, this effect is an additional 0.08 percent found in line 4 of table F-7.[36]

Differences in the definition of self-employment income make necessary a warning. JGF define self-employment income to be the difference between the proprietor's earnings less the return to capital used by the proprietor. On an hourly basis, this residual is generally a small fraction of the earnings of comparable employees. Denison used income from both capital and labor as self-employment income and thus assigned proprietors a much larger hourly wage than used by JGF. So, conceptual differences in the earnings of proprietors leads to some uncertainty concerning the magnitude of the effect of switching from the earnings weights of all workers to the earnings weights of employees only on the JGF measures.

JGF's measures also resemble Denison's in their inability to account for variations in the workweek by educational attainment. The matrix of hourly compensation is the product of several matrices all of which are estimated by JGF using the RAS method. The marginal distributions used to create these matrices imply how the distributions vary by traits. When no marginal distribution of cross-classified traits is available, the RAS method treats the distributions of traits as if they are independent of one another. JGF have marginal distributions of education in their employment and annual compensation per person matrices, but not in their hours per week and jobs per person matrices. As a result, JGF's estimates of hourly compensation are constructed on the assumption that hours per week are the same for all levels of educational attainment. Using the estimate in table F-7 for Denison, including variations in the workweek by education would lower JGF's labor composition growth by as much as 0.06 percent annually. In addition, JGF convert compensation per person to compensation per job. This requires dividing the matrix of compensation per person by a corresponding matrix of the number of jobs per person which JGF define as the number of weeks paid divided by 52. The jobs per person matrix is also generated using the RAS method without a marginal distribution for education. Since the number of weeks paid per year increases with educational attainment, accounting for variations in the number of jobs per person by education would further lower JGF's measures.

Though the figures above are only rough estimates, it appears that at least half of the difference between BLS and both Denison and JGF (the contribution of age, sex, and education only, bottom line of table F-5) can be accounted for by differences in the method of calculating earnings weights.

Summary

The BLS measure of labor composition grows 0.23 percent annually from 1948-79. Labor composition measures of Denison and of Jorgenson, Gollop, and Fraumeni (JGF) increase more than twice as fast over the same period. BLS measures the effect of changes in the work experience, education, and sex composition of the work force. Denison measured the effect of changes in age, education, sex, and labor force groups, and JGF cross-classified hours and earnings by age, education, sex, class of worker, and occupation.

Once these measures are reduced to similar concepts of education, age or work experience, and sex, Denison's and JGF's measures grow 0.1-0.2 percent faster than the BLS measures. At least half of this remaining difference can be accounted for by Denison's and JGF's inclusion of proprietors in the development of their earnings weights, Denison's use of full-time equivalent earnings in measuring the education earnings weights, and JGF's inability to account for variations in the number of hours per week and weeks paid per year by educational attainment.

[36] However, self-employed earnings rise faster with age than for employees. Omitting proprietors from the earnings weights would narrow the earnings differentials by age. Since the average age of the work force declines over the period of these studies, the effect of omitting proprietors from the earnings weights would dampen the difference in the growth rates of labor composition. Consequently, the effect of switching from an average earnings weight for all workers to employees would be less than the 0.08-percent estimate.

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Appendix G. Sensitivity Analysis of the Labor Composition Measures

The labor composition measures rely on both the hours and the associated hourly price for each identified type of worker. This appendix investigates the sensitivity of the labor composition measures to labor prices and the distribution of hours. This analysis can help identify which assumptions are critical to the measures and also indicate the robustness of the measures. Since the same labor prices are used in both sectors, the sensitivity of the two sectors to the myriad of assumptions is nearly identical. The analysis only addresses the private nonfarm business sector.

In the first section, the sensitivity of the labor composition measures to the price of labor is explored. Alternative schooling and work experience parameters are substituted, and labor composition is recalculated. In addition, three alternative labor prices are assigned to women, and the impact on labor composition is determined. In the next section, labor composition is measured using an altered distribution of hours. The sensitivity of the labor composition measures to the model of work experience is then examined by comparing reduced form and structural wage models. The final section is devoted to the effect of sampling error in the parameters of both the wage and experience models and the hours matrix.

The price of labor

The discussion in appendix A indicated that the wage model, and therefore the price of labor, are subject to alternative interpretations. These alternatives can be represented as a modification of the schooling and work experience parameters. Although the magnitude of alternative parameters is uncertain, this section quantifies the effect of a variety of alternative parameters on the labor composition measures.

Table G-1 contains estimates of the annual average labor composition growth rates for the period 1948-89 simulated under a number of assumptions. The parameters of the wage model are modified to reflect alternative views of how education and work experience affect the marginal product of labor. Simulations are done using labor prices generated from hypothetical parameters of the wage model. These hypothetical parameters are obtained by scaling the actual parameters (multiplication by a constant). The incremental return to additional schooling or experience is altered proportionately for all persons. For example, if each of the 12 schooling parameters (6 for men and 6 for women) are halved, the increase in wages due to additional schooling is also halved at each level of schooling. The four work experience parameters (2 for men and 2 for women) can be treated analogously or all 16 education and experience parameters can be altered simultaneously.

Table G-1. The effect of hypothetical education and work experience parameters on the annual average rate of growth of labor composition, 1948-89

Scale of parameters		Simulated labor composition	Effect of scaled parameters on labor composition
Education	Experience		
Sensitivity to education parameters			
1.00	1.00	[1]0.27	-
.90	1.00	.23	-0.04
.75	1.00	.17	-.10
.50	1.00	.07	-.20
1.25	1.00	.37	.10
Sensitivity to experience parameters			
1.00	1.00	[1]0.27	-
1.00	.90	.28	0.01
1.00	.75	.29	.02
1.00	.50	.31	.04
1.00	1.25	.25	-.02
Sensitivity to education and experience parameters			
1.00	1.00	[1]0.27	-
.90	.90	.24	-0.03
.75	.75	.19	-.08
.75	1.25	.15	.12
1.25	1.25	.35	.08
1.25	.75	.39	.12

[1] Actual labor composition measure.

The first section of table G-1 demonstrates a modest sensitivity of labor composition growth to the education parameters. A scale of 1.00 represents the actual parameters used throughout this study. A scale value lower than 1 narrows earnings differences by educational attainment, and scale values greater than 1 magnify this differential. The table indicates that labor composition growth increases with scale. Since earnings increase with educational attainment, increasing educational attainment of the work force adds to labor composition growth. Increasing the return to education or scale of the parameters would then lead to faster labor composition growth. The signaling and screening models or the correlation between education and ability, discussed in appendixes A and B, suggests that the marginal product of schooling is less than the measured earnings differentials, and schooling parameters would be smaller. However, even if the parameters were reduced 10 or 25 percent, annual labor composition growth would be reduced at most 0.1 percentage point. To place this scale in perspective, a 50-percent reduction in the parameters is well beyond statistical chance and would imply that screening (sometimes known as a sheepskin effect) was equally as important a determinant of earnings as education. The effect on the measures of multifactor productivity would be even smaller.[1]

The second section of table G-1 demonstrates a weak sensitivity of labor composition growth to work experience. Again, a scale value lower than 1 narrows earnings differences by experience, and scale values greater than 1 magnify this differential. Earnings generally increase with experience, and so declining levels of work experience subtract from labor composition growth. Increasing the return to experience or scale of the parameters would then lead to even slower labor composition growth. As a result, labor composition growth is negatively related to the scale of the experience parameters. The weak sensitivity of the measure results principally from the small decreases in the average level of experience. Consequently, smaller experience parameters consistent with the implicit contracts literature would produce a small increase in the annual labor composition growth and a negligible effect on multifactor productivity.

The final section of table G-1 shows the effects of varying both education and experience parameters. In general, the effects are greater than those of varying the experience parameters alone, but less than those of varying just the education parameters. Even at the extremes of scale such as 0.75 or 1.25, the annual average labor composition growth would differ by only -0.08 percent. Only when the education parameters are reduced and the experience parameters increased (or the converse) can labor composition growth differ substantially from the actual measures.[2]

The earnings of men and women show marked differences. This is confirmed in appendix E by F-tests that demonstrate the parameters of men and women are statistically different. As a result of these tests, the parameters are permitted to differ between men and women in the earnings equations. These tests by themselves cannot indicate the cause of the earnings differences. If the earnings differences between men and women are predominantly due to differences in skills, the parameters differences reflect differences in marginal products. If wage discrimination is present, wages may not equal the value of a worker's marginal revenue product. In such cases, the wages of women are likely to be less than the value of their marginal products. Since men are assumed to be paid the value of their marginal product, the wage parameters of men can be used to determine the value of a woman's marginal product in the presence of discrimination.

To explore the sensitivity of the labor composition measures to the possibility of sex discrimination, the wages of women can be adjusted to remove possible sources of wage discrimination. Three alternatives are used which reflect increasing levels of wage discrimination. In the first alternative, women receive different payments from men for their skills acquired through education and work experience, but receive equal payment for all other unmeasured skills and traits.[3] In the second alternative, women receive different payments for their skills acquired though work experience, but otherwise receive equal payment for all skills and traits.[4] In the third alternative, men and women receive equal payment for all skills and traits.[5] The average annual labor composition growth for the period 1948-89 for each alternative is presented in table G-2.

Table G-2 indicates that the use of alternatives 1 and 2 yields negligible changes in labor composition growth. Only when all wage differentials are attributed to sex discrimination can the difference, 0.11 percent annually, be viewed as nonnegligible. In the intermediate cases of alternatives 1 and 2, labor composition grows 0.02-0.03 percent faster annually compared to the base case.

In these alternatives, the parameters for women are set to the corresponding parameters for men. If men are the principal beneficiary of sex discrimination, men's wages overstate male productivity. It would then have been more appropriate to estimate parameters based on a pooled sample of men and women. This would lower the parameters and wages for men and raise them for women. The resultant alternative measure of labor composition would not differ as greatly from the base case as the figures shown in alternative 3.

[1] Since labor compensation is approximately two-thirds of total income, the effect of reducing the education parameters is only two-thirds as great as the effect on labor composition. Multifactor productivity growth would be affected by 0.03-0.13 percent annually.

[2] The discussion in appendix E suggests that wage models which do not use measures of actual work experience lead to overstated education parameters and understated experience parameters. Consequently, such wage models would lead to substantially higher estimates of labor composition than found in this study. For more on this issue, see appendix F for a comparison of several studies of labor composition.

[3] The wage model has three components: A payment for education as measured by the schooling parameters, a payment for on-the-job training as measured by the work experience parameters, and a payment for all other unmeasured skills as measured by the adjusted intercept of the wage model. In the standard model, all these payments may differ by sex. Under alternative 1, only the first two payments may differ, and the female intercept is set to the male intercept. Note that the male intercept is not larger than the female intercept in all years.

[4] Under alternative 2, the intercept and the education parameters for women are set to the corresponding parameters for men. The experience parameters are unchanged. This alternative effectively assumes that all male/female wage differences result from discrimination except that women acquire less on-the-job training than men.

[5] Using alternative 3, all parameters for women are set to the corresponding values for men. This is the equivalent of assuming all wage differentials are the result of discrimination.

Table G-2. The sensitivity of the average annual labor composition growth rate to male/female differences in the price of labor, 1948-89

Item	Labor composition growth	Difference from base case
Base case: All male/female labor price differences permitted	0.27	–
Alternative 1: Education and experience male/female labor price differences permitted	.29	0.02
Alternative 2: Experience male/female labor price differences permitted	.30	.03
Alternative 3: No male/female labor price differences permitted	.38	.11

Distribution of hours of work

The actual distribution of hours worked by education and work experience may differ from the measured distribution for several reasons. The Current Population Survey and decennial censuses are both subject to sampling and nonsampling errors. In addition, the work experience equation may allocate hours of work to an inappropriate level of work experience due to stochastic error, selection bias, or incomplete specification of the experience equation. Regardless of the cause, the measurement of labor composition would be affected.

This section measures the sensitivity of labor composition to the distribution of hours. The hours matrix is divided into 7 educational and 72 experience groups for each sex, and so it is not possible to explore all possible alternatives. Instead, only two are examined. Generalization to other cases should be made with extreme caution.[6]

Table G-3 shows the sensitivity of labor composition to changes in the educational distribution. The estimates in the table are made by measuring the additional increase in labor composition that would have occurred if the base-year matrix of hours were left unchanged and if the distribution of hours were altered in the following year. Specifically, the hours of each type of high school graduate are reduced a fixed percentage, and these hours are added to the corresponding hours of college graduates.

The data presented are in two forms. The first column indicates that a 0.01-year increase in the average level of educational attainment results in an increase in the labor composition annual growth rates ranging from 0.03 to 0.09 percent. The second column indicates that a shift of 1 million hours of work from high school graduates to the corresponding college graduates increases the annual labor composition growth rate from 0.11 to 0.30 percent.

A shift of 1 million hours or .01 years of educational attainment is quite large. Standard errors are not available for the mean years of education. However, the standard error for the percentage of employed labor force who are college graduates is 0.1 according to the 1971 CPS. Assuming this estimate applies as well to the percentage of hours, a 95-percent confidence interval implies that college graduates worked between 7.5 and 7.9 percent of total hours (7.7 ± 0.2). A 1-million-hour shift would increase the percentage of hours worked by college graduates to 8.6 percent, and a .01 increase in educational attainment would increase it to 8.0 percent. As can be seen, both these shifts are outside the confidence interval and are highly unlikely to occur.

Table G-3. The sensitivity of the annual average growth rate of labor composition in private nonfarm business to changes in the distribution of high school and college graduates, selected years, 1949-89

Year	Change in annual labor composition growth rate	
	0.01-year increase in educational attainment (percent)	1-million-hour shift (percent)
1949	0.03	0.14
1950	.04	.17
1955	.06	.25
1960	.07	.30
1965	.06	.23
1970	.05	.19
1975	.03	.12
1980	.04	.11
1985	.07	.16
1989	.09	.19

The sensitivity of labor composition is influenced by two trends: The first trend is the proportion of total hours that are shifted. Other things equal, a larger fraction of hours shifted between groups increases the sensitivity of the labor composition measures. In the first column, the proportion of hours shifted remains approximately constant, and so this trend does not alter the senstivity of labor composition over the period. In the second column, the relative size of the shift in hours declines because a fixed number of hours are shifted in a growing work force. This leads to decreasing sensitivity of labor composition. The second trend is the generally increasing earnings differential between high school and college graduates. This trend increases the sensitivity of the labor composition measure to shifts in hours, especially in the 1950's and 1980's.[7] In the first column, the second trend

[6] Because the wages are nonlinear with respect to both schooling and experience, the sensitivity of the labor composition measures depends on which set of hours is altered. In general, the sensitivity will differ for each possible alternative. High school and college graduates are used because they make up large percentages of total hours. There is no level of experience which accounts for a large fraction of the matrix, but 10 and 5 years of experience are among the largest.

[7] An increasing earnings differential over time implies that substituting a low paid worker with a higher paid one will result in increasingly larger additions to output in latter years. Given the same number of hours worked before and after the substitution, the increase in output appears as an increasingly large labor composition effect.

dominates and so, on balance, labor composition is increasingly sensitive over the period to shifts in hours. In the second column, the importance of the trends is reversed, and labor composition is decreasingly sensitive to shifts in hours.

Table G-4 shows comparable labor composition sensitivities to the distribution of work experience. The estimates in the table are made by measuring the additional change in labor composition that would have occurred if the base-year matrix of hours were left unchanged and if the distribution of hours in the following year were altered. Specifically, the hours of each type of worker with 10 years of experience are reduced a fixed percentage, and these hours are added to the corresponding hours of workers with 5 years of experience. There is not sufficient information to estimate the likelihood that these shifts are within the bounds of sampling error. However, it seems reasonable to believe that a 1-million-hour shift has approximately the same likelihood of falling within the confidence intervals whether shifting the distribution of education or work experience. A 0.01-percent change in the average level of work experience requires 0.2 percent of total hours to be shifted.

Once again, the data in table G-4 are presented in two forms. The first column illustrates the sensitivity of the annual labor composition growth rates to a 0.01-year decrease in the average level of experience. The sensitivity of the labor composition measures to changes in work experience is less than the sensitivity to changes in education. This largely reflects the smaller earnings differential between workers with 5 and 10 years of experience than between high school and college graduates. The sensitivity to a 0.01-year decrease in work experience gradually increases during the period and ranges from -0.02 to -0.04 percent. The second column shifts 1 million hours from each type of worker with 10 years of experience to the corresponding worker with 5 years of experience. This results in a mildly decreasing sensitivity which ranges from -0.09 to -0.15 percent.

Table G-4. The sensitivity of the average annual growth rate of labor composition in private nonfarm business to changes in the distribution of workers with 5 and 10 years of work experience, selected years, 1949-89

Year	Change in annual labor composition growth rate	
	0.01-year decrease in average experience (percent)	1-million-hour shift (percent)
1949	-0.02	-0.14
1950	-.03	-.15
1955	-.03	-.15
1960	-.03	-.14
1965	-.03	-.13
1970	-.03	-.14
1975	-.03	-.15
1980	-.03	-.11
1985	-.04	-.11
1989	-.04	-.09

As with the education distribution, the same two trends influence the sensitivity to experience. As with education, the earnings differential between the two experience levels increases consistently over the period. Both columns indicate an increasing sensitivity of labor composition to shifts in the distribution of work experience. The second column is nearly constant because the growing earnings differential nearly offsets the diminishing impact of a shift of a fixed number of hours in a growing work force.

The substantial annual changes in the growth rate of labor composition are the result of a large redistribution of hours. Even if such a misclassification had occurred, its impact on the labor composition growth rate for the entire period would be slight. That is, a 0.3-percent error in labor composition for a single year would add less than 0.01 percent to the average annual growth rate of the 1948-89 period. The point is that a biased distribution in an hours matrix for a single year is not large enough to alter the interpretation of the role of labor composition in productivity change. There would have to be a systematic, persistent, and growing bias in the measurement of the distribution of hours to alter the findings of this report. At most, this sensitivity analysis should caution the reader not to infer too much from changes in labor composition over very short time periods.

Labor composition growth using a reduced form model

The experience and wage models reflect a structural relationship between the determinants of work experience and their impact on earnings. An important example is the role of education which affects earnings both directly and indirectly. Education affects earnings directly as a source of skills developed in school. Education also changes the likelihood that a person will be in the labor market and acquire skills through work experience. Separate equations for work experience and earnings are needed to distinguish between these two effects, but they are not needed to measure labor composition. If the sole purpose of estimation is to develop a measure of hourly earnings for each type of worker, a reduced form model must work just as well if not better. Recall that earnings are a function of education and work experience, and work experience is a function of education, age, and other traits. A model which directly estimates earnings as a function of education, age, and other traits must provide at least as good a measure of hourly earnings (that is, a more efficient estimate) since there are fewer restrictions on the parameters.

In addition, the experience model is estimated using data for 1973 only. If the parameters or the structure of the experience model change over time, experience measures based on a model at a single point in time are likely to be increasingly biased in years distant from 1973. A reduced form model implicitly allows the parameters of the experience equation to change.

The empirical importance of these issues can be tested by estimating reduced form models and using the resultant estimates of hourly earnings to measure labor composition. A large difference in labor composition growth rates may indicate that the experience equation has not been stable over the last 40 years and that a single experience equation is insufficient.

Because of the enormity of the task of re-estimating all the wage models, a reduced form wage model is estimated for only 1949, 1959, 1967, 1972, 1977, 1982, and 1987. Hours matrices classified by experience are not useful in this context. A preliminary step to estimating labor composition produced hours matrices classified by education, age, marital status, number of children, and sex. These matrices are available for all years, 1948-89.

Using these data and the reduced form coefficients, labor composition change between 1949 and 1987 is estimated.[8] To isolate the effect of using a structural model rather than a reduced form model, labor composition using the structural coefficients is recalculated using only the parameters for the same 7 years.

The reduced form model contains all the variables included in either the experience or the wage equation. Furthermore, the wage equation contains the square of work experience, and so the cross-product of all the terms in the experience equation is also included in the reduced form model. (One obtains the cross-product by multiplying each term in the experience equation by every other term in the experience equation). The resulting wage equation for women has about 140 variables.[9] Table G-5 indicates that with the possible exception of the 1959-67 period labor composition grew at the same rate using either the reduced form or structural wage model parameter estimates.[10]

[8] Coefficients for any given year are based on the last previous estimated parameters, for example, 1950-58 use the 1949 parameters.

[9] Estimated parameters are not shown. On average the unadjusted R^2 increases 0.021 for women and 0.004 for men. These differences are significant at the 95-percent confidence level for both men and women in all years.

[10] A high degree of multicollinearity leads to a few implausible parameter estimates and impossible wage rates for some women. Certain groups of women (almost entirely never married women with four or more children) received either an infinite or zero wage. This occurs because there were no such women in the sample used to estimate the wage equations. If there had been, minimizing least squares would have forced a more reasonable estimate. There were a few such women in the hours matrix because the hours matrix includes unpaid family workers and proprietors. Assigning an infinite wage to these women effectively makes the compensation shares of other workers zero and reduces the labor composition measures to just a few cells because the compensation shares of these women dominate the measure. If these wages were included, labor composition growth would be distorted for the 1949-59 and the 1967-72 periods. In the other periods, impossible wage rates pose no problem because the cells are empty.

The impossible wage rates suggest that including the complete set of variables in the reduced form model is not a practical alternative. However, the point of this exercise is to determine if allowing the experience equation parameters to implicitly change over time affects the labor composition measures. Rather than stop at this point, very high wage rates are constrained to about $150 per hour and very low wage rates to about $0.006 per hour. This seems in keeping with the spirit of the reduced form wage equations which indicated rightly or not that these people were paid a very high or low wage. Very few women (never more than 0.4 percent) and no men were affected by this constraint.

Table G-5. Labor composition change in private business using reduced form and structural wage model parameter estimates, 1949-87

(Percent change)

Period	Reduced form	Structural model	Difference
1949-59	4.41	4.74	-0.33
1959-67	3.98	2.40	1.58
1967-72	.19	.31	-.12
1972-77	-.74	-.54	-.20
1977-82	2.15	2.09	.06
1982-87	1.52	1.46	.06
1949-87	11.96	10.84	1.12
Average annual growth rate, 1949-87	.30	.27	.03

Table G-5 indicates that in 5 out of the 6 periods the two methods yield virtually identical labor composition measures. While the reduced form model yields slightly faster growth for the 1959-67 period, the clear pattern of the results indicates that the structural parameter estimates yield labor composition measures very similar to those of the reduced form. Assuming the reduced form is generally capturing shifts in the experience parameters, then the structural model also is capturing shifting experience parameters through annual estimation of the wage equation.

This test does not disprove the possibility that the rate at which women acquire experience has changed over time. Since using a fixed 1973 experience equation or a reduced form wage model yields the same labor composition growth rates, two conclusions are possible. First, annual wage equations with imperfectly measured experience may be capturing shifts in the quantity of experience through shifts in the parameters that measure the return to experience. If so, the finding of increasing returns to experience over the period should be qualified. In this case, possible underestimates of experience are offset by overestimates of the return to experience. If experience is increasingly underestimated the further one goes from the 1973 model, the return to experience will be increasingly overestimated as well. Under this scenario, it is much less clear that the return to experience has been increasing. However, the wage rate for each type of worker is satisfactorily measured because as table G-5 indicates the two biases offset each other.

A second possible explanation for the equivalence of the reduced form and structural models is that changes in the level of work experience are largely determined by factors included in the wage model. That is, changes in the fertility rate, marital status, educational attainment, and age were the primary causes of changing levels of work experience for women rather than secular increases in labor force participation.

Sampling error

To this point, the discussion of the sensitivity analysis has focused on alternative parameters and distributions of hours of work. In this section, all these considerations are put aside. That is, the method of measuring labor composition is not directly considered. Instead, the only source of uncertainty in the measurement of labor composition considered in this section arises from the unavoidable sampling error in the hours matrices and the stochastic error of the parameter estimates.

Using the known stochastic distribution of the parameters and the sampling distribution of the hours of work, the parameters and the hours in each cell are randomly varied, and labor composition is recalculated.[11] For each year from 1968 to 1989, this process is repeated 200 times, and a sample standard deviation for labor composition is calculated.[12] The results are presented in table G-6.

Labor composition in the private nonfarm business sector grew 0.26 percent annually between 1968 and 1989, and the standard deviations, on average, are only slightly less than the annual labor composition growth rates. This suggests that random error in both the parameters and the hours matrices in most years can account for the annual increase in labor composition. However, labor composition growth over longer periods cannot be explained by random error and represents a real change in the skills of workers. For example, annual labor composition growth between 1982 and 1985 never exceeded two standard deviations in any year, but the labor composition index in 1985 is significantly different from the 1982 index. Therefore, use caution in analyzing and interpreting *annual* changes in labor composition and its impact on multifactor productivity. Changes in labor composition and its effects on labor input and multifactor productivity should be viewed as a *long term* phenomenon.

[11] The standard errors of the parameters of both the wage and experience model are known. Given the large sample size, the parameters can be assumed to be normally distributed with a mean and standard deviation measured by the estimated parameter and the parameter's standard error. The standard error of an individual cell of the hours matrix can be derived from published tables of *Employment and Earnings* which describe sampling error in the CPS. The percentage of total hours worked by any group of workers is assumed to have the observed mean and variance = p*(100-p)/N where p is the observed mean percentage and N is the sample size. The percentage of total hours worked by each group is assumed to vary normally although a binomial distribution would be have been more appropriate. Given the large sample size, a normal approximation is suitable. Whenever a random event produced a negative number of hours worked, the hours were set to an negligibly small but positive number.

[12] Sampling error is not examined prior to 1968 because the hours matrices were generated with the aid of a multiproportional interpolation method (RAS). The estimation error of RAS combined with the CPS sampling error is beyond this study. Since the RAS method must introduce some error, the standard errors of labor composition measures prior to 1968 are larger than those presented in table G-6. See appendix D for a discussion of the use of the RAS method.

Table G-6. Labor composition growth in private nonfarm business and its estimated standard deviation based on Monte Carlo simulation, 1968-89

Year	Labor composition (in percent) (1)	Standard deviation (2)	Ratio (1)/(2)
1968	-0.30	0.18	-1.67
1969	.43	.17	[1]2.53
1970	.41	.17	[1]2.41
1971	-.28	.20	-1.40
1972	-.04	.20	-.20
1973	-.21	.21	-1.00
1974	.68	.20	[1]3.40
1975	.00	.21	.00
1976	-.29	.22	-1.32
1977	.05	.18	.28
1978	.06	.19	.32
1979	-.32	.20	-1.60
1980	.33	.20	1.65
1981	.76	.18	[1]4.22
1982	1.07	.20	[1]5.35
1983	.37	.19	1.95
1984	.06	.16	.38
1985	.25	.18	1.39
1986	.58	.20	[1]2.90
1987	.20	.19	1.05
1988	.80	.19	[1]4.21
1989	.48	.19	[1]2.53

[1] Labor composition growth is significantly different from zero at the 95-percent confidence level (t-ratio > 1.96).

References

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Appendix H. Partial Indexes of Labor Composition

Since the path breaking work of Solow (1957) showed that output growth could be divided into the separate contributions of labor, capital, and productivity, there has been a natural tendency to decompose the contributions into finer and finer detail. Denison (1985), Jorgenson, Gollop, and Fraumeni (1987), and BLS have measured the contribution of shifts in the composition of the work force to labor input and productivity.

Exact measures of the separate contributions of each trait require a set of highly unlikely assumptions. An hour of work must be divisible into separate service flows for each trait. Furthermore, the service flows must be linear and additive. This imposes a linear structure without interaction terms on the relationship between wages and worker traits. Even though a linear wage model can be consistent with a hedonic wage model and Rosen (1974) has shown it must be the wage model if quantities of traits are perfectly divisible, most literature indicates that a semilog wage model is a more appropriate specification. Consequently, there is not likely to be any exact decomposition of labor input into the contribution of each trait.

Instead, this appendix describes a method for measuring the effect of shifts in one worker trait at a time holding the other two traits fixed. This differs slightly from an exact decomposition of labor composition because the wage model developed in appendix E is semilog and not linear. In the next section, the additional assumptions required for unambiguous measures of the contribution of each trait to labor input and composition are discussed. It will be seen that the methodology developed later in the section does not satisfy the restrictions because a linear wage model is an improbable specification of the relationship between wages and worker traits.[1] In the final section, measures of the main and interaction effects of education, work experience, and sex composition are presented and discussed.

Labor composition

By assuming competitive factor markets and a production function which exhibits constant returns to scale and separability of labor from other inputs, one can obtain a measure of labor input, L, which is determined by the hours of many different types of workers.[2] For the purposes of this appendix, the hours of workers are classified over three traits: Education, experience, and sex. A worker of educational attainment, e, work experience, x, and sex, s, supplies hours h_{exs}. The aggregator function for all hours over the entire range of possible types of workers is:

$$(1)\ L = L(h_{111}, \ldots, h_{exs}, \ldots, h_{xyz})$$

Taking the log of labor input and then its derivative with respect to time yields an expression for the growth rate of labor input.

$$(2)\ d \ln L/dt = \Sigma_{exs}\, (\partial \ln L \,/\, \partial \ln h_{exs}) * d \ln h_{exs}/dt$$

Equation 2 states that the growth rate of aggregate labor input is a sum of the growth rates of the hours of a specific type of worker weighted by the elasticities of aggregate labor input with respect to the hours of that specific type of worker. Competitive factor markets lead firms to equate each elasticity with its share of labor costs, and equation 2 can be rewritten as:

$$(3)\quad d \ln L/dt = \Sigma_{exs}\, v_{exs} * d \ln h_{exs}/dt$$

where v_{exs} is the share of total compensation paid to a given type of labor paid an hourly wage rate w_{exs}.

$$(4)\quad v_{exs} = w_{exs} * h_{exs} \,/\, \Sigma_{exs}\, (w_{exs} * h_{exs}).$$

The growth rate of labor composition, LC, is defined as the difference between the growth rate of labor input and the growth rate of total hours, d ln H/dt.

$$(5)\quad d \ln LC/dt = d \ln L/dt - d \ln H/dt$$
$$= \Sigma_{exs}\, v_{exs} * d \ln h_{exs}/dt - d \ln H/dt$$

As already seen in appendix A, equation 5 is the starting point for the framework for measuring labor composition and its components. If labor composition can be decomposed into separate contributions from each trait, implicitly, labor services flow simultaneously from each of the three traits. The labor input of a worker, l_{exs}, is a function of three services and the number of hours, h_{exs}. For a worker with a given education, e, work experience, x, and sex, s, the input for each type of worker is:

[1] It has been shown by Rosenblum, Dean, Jablonski, and Kunze in "Measuring Components of Labor Composition Changes," presented at the American Economic Association meetings, December 1990, that the methods used by Denison and Jorgenson, Gollop, and Fraumeni also fail to satisfy an exact decomposition.

[2] See E. Berndt and L. Christensen, "Testing for the Existence of a Consistent Index of Labor Inputs," *American Economic Review*, June 1974, for a discussion of separable inputs.

$$(6)\ l_{exs} = l\,(e*h_{exs}, x*h_{exs}, s*h_{exs}) = l(e, x, s)*h_{exs}$$

Equation 6 makes explicit the notion that the productive value of an hour depends on a worker's characteristics. This form is multiplicative in hours to maintain linear homogeneity in aggregate labor input. That is, a doubling of the hours of each type of worker doubles labor input. Using equation 6, aggregate labor input (equation 3) can be expressed as:[3]

$$(7)\ L = L\,(\ldots, l\,(e * h_{exs}, x * h_{exs}, s * h_{exs}), \ldots)$$

Taking the log and then the derivative of equation 7,

$$\begin{aligned}(8)\ & d \ln L/dt = \\ & \Sigma_{exs}\ \partial \ln L/\partial \ln l_{exs} * \partial \ln l_{exs}/\partial \ln e{*}h_{exs} * d \ln e{*}\ h_{exs}/dt \\ & +\Sigma_{exs}\ \partial \ln L/\partial \ln l_{exs} * \partial \ln l_{exs}/\partial \ln x{*}h_{exs} * d \ln x{*}\ h_{exs}/dt \\ & +\Sigma_{exs}\ \partial \ln L/\partial \ln l_{exs} * \partial \ln l_{exs}/\partial \ln s{*}h_{exs} * d \ln s{*}\ h_{exs}/dt\end{aligned}$$

Again assuming factor markets are competitive, the labor aggregate is linear homogeneous with respect to the labor inputs, and the quantity of services per unit of each trait is constant over time, equation 8 simplifies to:[4]

$$\begin{aligned}(9)\quad & d \ln L/dt = \\ & \Sigma_{exs}\,(PE_{exs} * e * h_{exs}/w * H) * d \ln h_{exs}/dt \\ & +\Sigma_{exs}\,(PX_{exs} * x * h_{exs}/w * H) * d \ln h_{exs}/dt \\ & +\Sigma_{exs}\,(PS_{exs} * s * h_{exs}/w * H) * d \ln h_{exs}/dt\end{aligned}$$

where w is the aggregate average wage rate, H is total hours, and PE, PX, and PS are the unit trait prices of workers' education, experience, and sex. The unit price for each level of education, for example, is the derivative of the production function with respect to the quantity of a trait ($\partial Q/\partial L * \partial L/\partial l_{exs} * \partial l_{exs}/\partial\,(e{*}h_{exs})$).

There is no requirement that the price of a given level of a trait (for example, college graduate) must be the same for all types of workers (such as young and older workers). However, if each college graduate has a unique price for her or his schooling, and each worker with 5 years of experience has a unique price for his or her experience, and each person has a unique price for her or his residual traits, then it is not possible to distinguish among payments for education, experience, and residual traits.[5] Consequently, it is assumed that a single price exists for each level of each trait. Note that some differences in the price of the same level of a trait are permissible and measurable. However, those differences which are not purely ad hoc cause the measurement of the separate contributions of each trait to break down as shown below. The point is that some restrictions on the price of each trait are required to measure the contribution of each trait to labor input. Using this assumption, equation 9 can be rewritten.

$$\begin{aligned}(10)\quad & d \ln L/dt = \\ & \Sigma_{exs}\,(PE_e * e * h_{exs} / w * H) * d \ln h_{exs}/dt \\ & + \Sigma_{exs}\,(PX_x * x * h_{exs} / w * H) * d \ln h_{exs}/dt \\ & + \Sigma_{exs}\,(PS_s * s * h_{exs} / w * H) * d \ln h_{exs}/dt \\ & = s_e * \Sigma_{exs}\,(PE_e{*}e * h_{exs} / \Sigma_{exs}\, PE_e{*}e{*}h_{exs}) * d \ln h_{exs}/dt \\ & + sx * \Sigma_{exs}\,(PX_x{*}x * h_{exs} / \Sigma_{exs}\, PX_x{*}x{*}h_{exs}) * d \ln h_{exs}/dt \\ & + ss * \Sigma_{exs}\,(PS_s{*}s * h_{exs} / \Sigma_{exs}\, PS_s{*}s{*}h_{exs}) * d \ln h_{exs}/dt\end{aligned}$$

where s_e (= $\Sigma_{exs} PE_e * e * h_{exs}/w{*}H$), s_x, and s_s are the share of total compensation paid to each trait. Equation 10 provides a method for accounting for the contribution of traits to labor input growth in which the last three lines measure the separate contributions of each trait. Finally, the contribution of each trait to labor composition growth is derived by subtracting the growth rate of total hours from the growth rate of labor input. Since the payment shares, s_e, s_x, and s_s, must sum to one, the growth rate of labor composition can be expressed as:

$$\begin{aligned}(11)\quad & d \ln LC/dt = d \ln L/dt - d \ln H/dt \\ & = s_e * \Sigma_{exs}\,({}_ep{*}h_{exs} / \Sigma_{exs}\ {}_ep{*}h_{exs}) * (d \ln h_{exs}/dt - d \ln H/dt) \\ & + s_x * \Sigma_{exs}\,({}_xp{*}h_{exs} / \Sigma_{exs}\ {}_xp{*}h_{exs}) * (d \ln h_{exs}/dt - d \ln H/dt) \\ & + s_s * \Sigma_{exs}\,({}_sp{*}h_{exs} / \Sigma_{exs}\ {}_sp{*}h_{exs}) * (d \ln h_{exs}/dt - d \ln H/dt)\end{aligned}$$

where ${}_ep$ (= $PE_e * e$), ${}_xp$, and ${}_sp$ are the payments to education, experience, and residual traits.

Partial indexes

Attempts at a decomposition of labor composition into the contribution of each characteristic have essentially relied on developing a price for each trait. This new price is then substituted for a worker's compensation to develop new share weights in the formula for labor composition. A closer look at equation 11 reveals that it is simply a restatement of the definition of labor composition (equation 5). The sole difference is that the compensation share weight has been divided into three components representing the separate contribution of each trait.

It follows from equating equations 5 and 11, both equations for aggregate labor composition change, that the sum of the trait payments (${}_ep + {}_xp + {}_sp$) must equal the wage rate for each type of worker. This would impose a linear form on the wage model. This restriction is severe, but it is required to yield an

[3] Welch (1970) also derives a production function which depends on the education of workers. Welch's model does not consider experience and other traits as inputs. As such, Welch uses the quantity of education as the labor input instead of hours, but this model uses education as a factor which affects the rate of labor services per hour.

[4] Specifically, $de_j/dt = dx_j/dt = ds_j/dt = 0$ implies that the services per unit of each of the three traits does not change. Equation 8 indicates the weighted growth rate of hours and the weighted growth rate of service per unit (trait quality) are additive when measuring labor input. The measurement of these quality changes are beyond this study. See Denison, *Accounting for U.S. Economic Growth, 1929-69*, appendix I and J. Bishop, "Is the Test Score Decline Responsible for the Productivity Decline?," *American Economic Review*, March 1989, for a discussion of education quality changes in the United States.

[5] If the price of each level of each trait is unique, the payments for the education, experience, and residual traits of a given type of worker are perfectly collinear. In a regression, perfect collinearity prohibits separate estimates for each payment.

exact decomposition. Most of the work in the human capital wage model suggests that the log of wages is linearly related to a set of characteristics. Tests of the semilog versus the linear model (Heckman and Polachek, 1974) indicate that the semilog form provides a superior fit of the data. Consequently, if the true wage model is semilog or any nonlinear form, no exact decomposition of labor composition growth into components is possible.

The model described in appendix E is not a linear wage model. Instead it uses a semilog model to measure trait payments and models the log of earnings as a function of education and age (or work experience). As indicated above, only a linear wage model can provide an exact decomposition of labor input. Nevertheless, a semilog wage model is particularly well suited to the task of decomposing labor composition because the trait payments are multiplicative and can easily be held constant.

Maintaining the previous assumptions, substitution of a semilog wage model into equation 3 yields a different version of equation 10 in which separate contributions for each trait and separate trait payments do not exist.

(12) $d \ln L/dt =$
$\Sigma_{exs}\ ({}_{e}p * {}_{x}p * {}_{s}p * h_{exs} /(\Sigma_{exs}\ {}_{e}p * {}_{x}p * {}_{s}p * h_{exs}) * d \ln h_{exs} /dt)$

The strategy here is to substitute a new share weight for the compensation share weight in the formula for labor composition. This is accomplished by developing an estimate of the payment for a single trait. By holding the payment for two traits constant, changes in the distribution of these traits cannot affect labor input. To see this, substitute a pair of constants for two of the payments in the share weight of equation 12. These appear as multiplicative constants in both the numerator and denominator and sc cancel. Using the same nomenclature as Barger (1971) and Jorgenson, Gollop, and Fraumeni (1987), the first order partial index, I, of the effect of changes in the distribution of hours by education holding experience and sex composition constant is:

Education
(13a) $I_e = \Sigma_{exs}\ {}_{e}p * h_{exs} /((\Sigma_{exs}\ {}_{e}p * h_{exs}) * d \ln h_{exs} /dt)$

Similarly, first order partial indexes for work experience and sex composition can be defined.

Work experience
(13b) $I_x = \Sigma_{exs}\ {}_{x}p * h_{exs} /((\Sigma_{exs}\ {}_{x}p * h_{exs}) * d \ln h_{exs} /dt)$

Sex composition
(13c) $I_s = \Sigma_{exs}\ {}_{s}p * h_{exs} /((\Sigma_{exs}\ {}_{s}p * h_{exs}) * d \ln h_{exs} /dt)$

It is also important to note that this is not a "true" measure of the contribution of each trait to labor input. If separate prices existed that satisfy both the linear and semilog wage model (and there are none), equation 10 would measure the contribution of a trait to labor input along the production frontier. Equation 12 shows the effect of shifts in the distribution of education on labor input holding experience and sex composition fixed.

Second order partial indexes can be defined analogously to the first order partial indexes of equation 13. Second order indexes consider changes in the distribution of two traits simultaneously and hold only one trait constant.

Education-experience partial index
(14a) $I_{ex} = \Sigma_{exs}\ {}_{ex}p * h_{exs}/(\Sigma_{exs}\ {}_{ex}p * h_{exs}) * d \ln h_{exs} /dt)$
Education-sex composition partial index
(14b) $I_{es} = \Sigma_{exs}\ {}_{es}p * h_{exs}/(\Sigma_{exs}\ {}_{es}p * h_{exs}) * d \ln h_{exs} /dt)$
Experience-sex composition partial index
(14c) $I_{xs} = \Sigma_{exs}\ {}_{xs}p * h_{exs}/(\Sigma_{exs}\ {}_{xs}p * h_{exs}) * d \ln h_{exs} /dt)$

A third order partial index could also be defined, but it considers changes in all three traits simultaneously, and it is identical to the labor input index.

All that remains is to determine a set of trait payments based on the wage model described in appendix E. In general terms, the log of the wage, W, is a function of a set of dummy variables to measure education and first and second order measures of a continuous work experience variable. Separate equations are estimated for men and women.

(15 men) $\ln W_{exs} = am_s + bm_e + cm * X - gm * X^2$
(15 women) $\ln W_{exs} = af_s + bf_e + cf * X - gf * X^2$

With the exception of work experience, the parameters directly represent the log of the trait payments. The terms bm_e and bf_e are the parameters representing payments to each level of education for men and women, respectively. A single pair of work experience parameters is estimated for each sex. The first order parameters for men and women are cm and cf, and the second order parameters are gm and gf. The result, $cm * X - gm * X^2$, represents the payment for the experience of a person with X years of work experience. The male and female returns to all other traits included in the wage equation are measured by am_s and af_s.

The payment to each trait is not the average earnings for each level of each trait, as used for example by Chinloy and Jorgenson, Gollop, and Fraumeni, but instead is determined by the estimated parameters of the model.[6] Because two traits are held constant and these constants cancel out of the calculation, they can be conveniently set to zero. As a result, the payment to a person for his or her education is the exponentiation of the education parameter, bm_e. The payments to each trait setting the payment to the other two traits to zero are:

[6] Estimated parameters from the wages model are preferred over average earnings to estimate the payment for each trait. Based on average earnings, the earnings differential by education reflects not only differences between the educational attainment of different groups of workers but also differences in work experience.

The average level of educational attainment has been rising over the past 40 years. As a result, the average college graduate is younger than the average high school graduate. Since older workers generally have more work experience than younger workers, the wage differential between high school and college graduates as measured by average earnings is too narrow to reflect just the difference in educational attainment. However, the wage differentials based on the estimated education parameters measure the wage differential holding experience and all other factors constant.

For a discussion of the differences between the two methods, see Rosenblum, Dean, Jablonski, and Kunze, "Measuring Components of Labor Composition Changes."

Payment for education

(16 men) $\quad {}_{e}p = \exp(bm_e)$

(16 women) $\quad {}_{e}p = \exp(bf_e)$

Payment for experience

(17 men) $\quad {}_{x}p = \exp(cm * X - gm * X^2)$

(17 women) $\quad {}_{x}p = \exp(cf * X - gf * X^2)$

Payment for residual traits (sex composition)

(18 men) $\quad {}_{s}p = \exp(am_s)$

(18 women) $\quad {}_{s}p = \exp(af_s)$

Equation 15 does not include interaction terms between traits so that a nonarbitrary payment can be measured for each trait.[7] Suppose an interaction between education and experience was present in the wage model. Such an interaction term is equivalent to permitting (in equation 10) the price of a given level of a trait to differ across a subset of workers. The trait payment for both education and experience would then be indeterminate because there is no unambiguous method for dividing the interaction term between payments to education and experience. Measures of the separate contribution of each trait in such a case would be highly uncertain and arbitrary.

The set of trait payments used in the second order partial index hold one trait constant. These payments are used to develop share weights for the second order partial indexes.

Education-experience payment

(19 men) $\quad {}_{ex}p = \exp(bm_e + cm * X - gm * X^2)$

(19 women) $\quad {}_{ex}p = \exp(bf_e + cf * X - gf * X^2)$

Education-sex composition payment

(20 men) $\quad {}_{es}p = \exp(bm_e + am_s)$

(20 women) $\quad {}_{es}p = \exp(bf_e + af_s)$

Experience-sex composition payment

(21 men) $\quad {}_{xs}p = \exp(cm * X - gm * X^2 + am_s)$

(21 women) $\quad {}_{xs}p = \exp(cf * X - gf * X^2 + af_s)$

The object of this exercise is to determine a set of main and interaction effects. The growth rate of the main effect, Q, is obtained by subtracting the growth rate of total hours, H, from the first order partial index, I. First order interaction effects are defined as the second order partial index less the appropriate two first order indexes and the growth rate of total hours.

[7] Permitting parameters for education and experience to differ between men and women is implicitly an interaction effect between sex composition and experience and education. These parameter differences are assumed to reflect differences in the quantity of education and training embodied in a year of school or work experience and not a difference in the return to these traits. It is possible that male/female differences in the return to education and training might also be due labor market discrimination or an interaction between the quantity of education or training and other unmeasured abilities. Therefore, I_e and I_x include shifts in the distribution of education or experience caused by a shift in hours between men and women.

Main effects

(22a) $\quad Q_e = I_e - d \ln H/dt$

(22b) $\quad Q_x = I_x - d \ln H/dt$

(22c) $\quad Q_s = I_s - d \ln H/dt$

First order interaction effects

(22d) $\quad Q_{es} = I_{es} - d \ln H/dt - Q_e - Q_s$

(22e) $\quad Q_{ex} = I_{ex} - d \ln H/dt - Q_e - Q_x$

(22f) $\quad Q_{xs} = I_{xs} - d \ln H/dt - Q_x - Q_s$

Finally, the second order interaction effect is defined to insure that the main effects and interaction effects sum to labor composition growth.

Second order interaction effects

(22g) $\quad Q_{exs} = d \ln LC/dt - Q_{es} - Q_{ex} - Q_{xs} - Q_e - Q_x - Q_s$

Contributions of education, experience, and sex composition

The contributions of education, experience, and sex composition to labor composition growth in private business are shown in table H-1. (The figures for private nonfarm business are very similar and are not shown.) Remember that these are estimates and do not represent an exact decomposition of labor composition growth.

For the entire period from 1948 to 1990, the main effect of education on labor composition was 0.41 percent per year. In contrast, the main effects of experience and sex composition were negative and much smaller in magnitude; the main experience effect was -0.10 percent, and the main sex composition effect was -0.03 percent. The main effects account for almost all of labor composition growth between 1948 and 1990. Most of the interaction effects in that period are zero, and they sum up to merely 0.02 percent.

Table H-1. Decomposition of labor composition growth in the private business sector, 1948-90
(Percent per year)

Composition effects	1948-90	1948-73	1973-79	1979-90
Labor composition growth	0.30	0.27	0.05	0.49
Main effects				
Education	.41	.40	.50	.38
Experience	-.10	-.11	-.44	.10
Sex composition	-.03	-.04	-.06	-.01
Interaction effects	.02	.01	.05	.02
First-order:				
Education-experience	.02	.01	.04	.02
Education-sex composition	.00	.00	.00	.00
Experience-sex composition	.00	.00	.01	.00
Second-order:				
Education-Experience-sex composition	.00	.00	.00	.00

NOTE: The sum of the effects may not equal labor composition growth due to rounding.

Looking at the subperiods in table H-1, it is clear that the main effect of experience varies far more than any of the other effects. From 1948 to 1973, the main experience effect contributes -0.11 percent to labor composition growth, and this effect plummets to -0.44 percent between 1973 and 1979. This drop in the main effect of experience coincides with the entrance of many members of the baby boom generation into the work force and with the steadily increasing participation of women in the labor market. During the latest period, 1979 to 1990, the experience effect swings up by more than half a percentage point, to 0.10 percent. Comparing movements in labor composition growth and in the main effects, it is evident that the variation in the main effect of experience explains almost all of the variation in labor composition growth.

It is important to note that the contribution of the sex composition effect in this study is quite small. The small sex composition effect together with a largely unchanging contribution of education suggests that any reduction in labor composition growth due to the increased participation of women is almost exclusively identified with a reduction in the average level or return to experience of the entire work force.

Virtually all of the interaction effects in table H-1 are negligible. Only the education-experience effect is different from zero in each period although it never exceeds 0.04 percent. Because the interaction effects are so small, the sum of the main effects calculated with the partial indexes is approximately equal to the growth rate of labor composition.

The contribution of the main effects to labor productivity is measured by multiplying them by labor's share of current dollar output (which has been about two-thirds on average). The contribution of education to labor productivity was approximately 0.3 in each of the periods under consideration. In the latest period, 1979 to 1990, the contribution of education almost equalled the contribution of labor composition (0.3 percent). Labor composition, in turn, accounted for close to 40 percent of the annual growth rate of output per hour, which was 0.8 percent. The annual contribution of experience to labor productivity was -0.1 percent during both the 1948-90 and 1948-73 periods. Between 1973 and 1979, the annual contribution of experience dropped to -0.3 percent, while in the most recent period experience made a positive contribution of 0.1 percent. In each period the annual contribution of sex composition to productivity growth was 0.0 percent.

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